7

Ontario Ministry of Labour
The Honourable Fern Guindon, Minister
R. D. Johnston, Deputy Minister

Research Branch
March 1973

Ontario

Union-Nonunion Wage Differentials: A Cross-Sectional Analysis

Ontario Ministry of Labour
The Honourable Fern Guindon, Minister
R. D. Johnston, Deputy Minister

Research Branch
March 1973

Union-Nonunion Wage Differentials: A Cross-Sectional Analysis

Prepared by
Gerald Frank Starr

FOREWORD

This study seeks to document the determinants and overall size of union-nonunion wage differentials. It has been carried out through a detailed examination of intra-occupational wage differentials in the manufacturing industries of Ontario in 1969. Besides being inherently controversial, this is a difficult area of research due to problems in isolating the union influence from other factors that affect wage rate levels. Yet, studies of this kind are essential if any progress is to be made in improving our understanding of the role unions play in the wage determination process. It is surprising that, although there have been a few broad enquiries into the state of industrial relations in the post World War II period, there is, to my knowledge, no other general analysis of union-nonunion wage differentials in Canada.

The study was undertaken while the author was a member of the Research Branch of the Ontario Ministry of Labour and has been accepted as a Ph.D thesis at the University of Toronto. As in many quantitative studies of labour market behaviour, extensive use has been made of standard linear regression analysis techniques. However, for those readers who are mainly concerned with the results rather than methodology, an effort has also been made to point out clearly the implications of each section of this study. In addition a short non-technical exposition of the main findings is available from the Research Branch. In both these documents the views expressed are strictly personal. It should be clearly understood that the Ministry of Labour is not responsible for the inferences made or errors in the analysis.

Various people have made major contributions toward the completion of this study. It was made possible by John Kinley, Director of the Research Branch, who continually supported the project during its rather long history. Besides offering needed prompting and encouragement, Professor Morley Gunderson acted as my sounding board throughout the various stages of the study. If all his suggestions had been followed, the document would have been much improved. The members of my thesis committee, Professors Arthur Kruger, Sheila Eastman and Gregory Jump, each made many helpful comments on early drafts. Professor N. Choudhry also helped identify and correct some errors in the final draft. Mrs. Pat Hayes carried out the programming work with remarkable efficiency and consistent good humour while Ms. Pearl Suin was subject to my totally unreasonable typing demands. I am deeply indebted to all these people as well as other members of the Research Branch who helped in the project. A special debt is owed to my wife and children who stoically bore the byproducts of my personal obsession with this work.

Gerald Starr

TABLE OF CONTENTS

LIST OF TABLES

LIST OF CHARTS

INTRODUCTION

Unions can be viewed as having three types of effects on wages. First, they may change the general level of money or real wages throughout the economy. Second, they may influence the rate at which money wages are increasing. Finally, they may alter the wages of unionized labour relative to nonunionized labour. Each of these effects is quite distinct. In this study the focus is on the last of these, that is, the union role in shaping the wage structure.

Unions are important labour market institutions whose behaviour and economic effects need to be well understood if public policy is to be formulated on an intelligent basis. Conceivably they can raise the wages of their members above what they would otherwise be. If this is so, not only will the distribution of income be changed but also the allocation of resources in the economy. If in the absence of unions labour markets approximate competitive conditions, at least in the long-run, then a positive union-nonunion wage differential implies a misallocation of resources. Total output will be smaller than potentially it might be. But the effects may be large or small and may vary dramatically between circumstances. As the income distribution and resource allocation impacts are two of the many factors to be considered in assessing public policy towards unions, it seems critical to have rather precise knowledge concerning the distribution of union effects on the wage structure. In addition, such knowledge might give some clue as to the union impact on general wage levels or rates of wage change. Although the exact relationships have not been established, it seems unlikely that a major role could be attributed to unionism in either area if union effects on the wage structure are found to be minimal.

Despite the apparent importance of these questions, the number of in-depth studies of union impact has not been great in recent years. There are a number of reasons for this. Probably the most important has been a marked shift of emphasis in research away from microeconomic studies of wage structures of the kind that were quite common during the 1950's. During the 1960's, most labour economists have been concerned with patterns of labour force participation and mobility, investment in human capital, the economics of discrimination, and macroeconomic aspects of the labour market. This shift undoubtedly reflects new policy concerns, and perhaps, as well, the belief that the basic framework for collective bargaining is not likely to be changed. Interest in the area of union impact may also have waned due to the apparent intractability of some of the analytical problems and the unavailability of a suitable data base.

In the absence of a well-developed literature, no strong consensus has emerged on the union role in shaping the wage structure. At one extreme,

there are those economists who side with public opinion and attribute to unions a major role in wage determination. Indeed, unions are characterized as powerful institutions who have virtually insulated their members from market restraints. At the other extreme, many economists seem to minimize the union influence, except perhaps for temporary disequilibrium effects. In their view, free entry by nonunion firms generally ensures that competitive market forces are pervasive. The image created is one of a society that tolerates much ritual and noise in the form of collective bargaining largely to preserve an illusion. But the evidence supporting either extreme position is scarce. The three latest Canadian textbooks on labour economics have been forced to take the same essentially agnostic view on the question of union impact. This is a sad comment on the rate of progress in analysis of a major issue.[1]

The quantitative estimates of union relative wage effects that have been made recently remain obscure and appear to have convinced very few. In a review of the current state of labour economics, Albert Rees has remarked:

> Although much good work has been done on the impact of the union, there are still large divergencies between the estimates made by different investigators. Some of these reflect differences in the data used and others reflect differences in methods. Reconciling these estimates will require more work in the years ahead, and more thought on such issues as the effect of the threat of organization on the wages offered by nonunion employers.[2]

This is the direction in which this study is aimed.

In Chapter I the recent literature is reviewed. The wide range of estimates of overall union-nonunion wage differentials is documented as well

---

[1] Sylvia Ostry and Mahamood A. Zaidi, Labour Economics in Canada, Vol. II of Labour Policy and Labour Economics in Canada (Second ed.; Toronto: Macmillan of Canada, 1972), pp. 310-314; J. Tait Montague, Labour Markets in Canada: Processes and Institutions (Scarborough Ontario: Prentice-Hall of Canada, 1970) passim, pp. 211-239; Stephen G. Peitchinis, Canadian Labour Economics: An Introductory Analysis (Toronto: McGraw-Hill Company of Canada Limited, 1970), pp. 311-312, 436-441.

[2] Albert Rees, "The Current State of Labour Economics," Paper delivered at the conference of Canadian Labour Economists held at Queen's University, Kingston, Ontario, Feb. 26-27, 1971, Reprint No. 16, Industrial Relations Centre, Queen's University (Kingston, Ontario: Industrial Relations Centre, Queen's University, 1971), p. 5.

as evidence on variations in these differentials. A potential source of bias in the estimates, rooted in the methodologies of studies using industry aggregated data, is discussed. In Chapter II, a unique disaggregated data base available for analyzing the Ontario experience is described. This data base gives occupational wage rate and union status information at the establishment level. The chapter also provides a theoretical framework suitable for explaining wage rates at the establishment level. This framework provides the rationale for all the regression equations used in this study. Results for unskilled male and female labour are then analyzed in Chapters III and IV respectively. In Chapter V, selected blue-collar occupations are considered in order to gain some notion of variations in union effects by skill level. Next, the direct and indirect effects of unionism on white-collar wages are considered. Chapter VII is by way of summary.

CHAPTER I

REVIEW OF THE LITERATURE

The year 1963 is a convenient point of departure for this review. It was in this year that Lewis published his exhaustive examination of the available studies.[1] With the aid of a well-developed conceptual framework, he analyzed and reworked many of the results from the previous studies and carried out considerable original empirical work. Based on both time series and cross-section evidence, he concluded that during the second half of the 1950's, the union-nonunion differential averaged 10 to 15 per cent. This estimate has since served as the benchmark against which subsequent estimates have been compared.

This review of the literature will only cover studies completed after Lewis's book was published. In addition, studies based on data from single industries are excluded because they do not yield the kinds of estimates being considered here. Also, only studies on the United States experience are considered since there are no comparable Canadian studies.

All the studies reviewed (except one) are more or less similar in that they use a multiple regression analysis of cross-section data. However, they differ in a number of important respects. Only five studies have the individual or the establishment as the unit of observation. All the others, due to a lack of more detailed data, use industry averages as variables in the regression equations. As there are special problems in interpreting the results when this approach is used, these studies are grouped together in the discussion. The underlying analytical models also vary considerably. A detailed review of these differences is not feasible within the scope of this chapter. However, to give the reader some indication of the factors being controlled for in the wage determination equations, all the variables included are mentioned. A final distinction is the scope of the studies. While all give overall estimates of union-nonunion wage differentials, some also provide evidence on inter-industry and inter-skill variations in union effectiveness. Such information is noted whenever it is available.

---

[1] H. G. Lewis, Unionism and Relative Wages in the United States (Chicago: The University of Chicago Press, 1963).

Studies Using Industry Aggregated Data

Fuchs[2]

Fuchs was primarily interested in explaining the low wages found in the service sector. In order to do this, he used data from the one in 1,000 sample of the U.S. 1960 Census of Population. The dependent variable was average hourly earnings in a three-digit industry. Variations in labour quality were controlled for by assigning to each industry an expected average hourly earnings rate calculated by assuming that each worker in the industry had an hourly earnings rate equal to the national rate for workers with the same colour, age, sex, and education characteristics. Other independent variables used were region, city size, city location, establishment size, employment growth, unemployment rate, average annual hours per employed male, self-employment income as a percentage of total earnings, and union members as a proportion of employees in the industry. Regressions were run across 138 non-agricultural industries as well as a subset of these called the "industrial sector." This latter group consisted of 81 industries in manufacturing, mining, construction, transportation, communication, public utilities and the postal service.

With the dependent variable in a log form, the coefficients for the variable indicating extent of union organization in the "all industry" and "industrial sector" samples were .21 and .25 respectively. The problem is to determine under what conditions these coefficients can be interpreted as measures of a union-nonunion wage differential.

The problem of interpretation arises because a relationship between average earnings in an industry and the extent of unionism may be the result of three underlying relations. First, the relationship may arise simply due to weighting. Well-organized industries have a high proportion of presumably high-paid union workers. Second, there may be a relationship between the average wages of union firms in an industry and the extent of union organization. Unions that can organize virtually all of an industry may be more successful in obtaining high wages than other unions. Finally, a positive or negative relationship might exist between the wages paid in unorganized plants and the extent of union organization in the industry. Interpreting a regression coefficient that relates overall average wages in an industry to the extent of union organization requires that assumptions be made about each of these three possibilities.

---

[2] Victor R. Fuchs, The Service Economy (New York: The Columbia University Press, 1968), pp. 127-59.

One interpretation, occasionally given,[3] is that the coefficients basically indicate differences in union effectiveness between industries. If it is assumed that in an industry union and nonunion wages are the same and that the union impact on wages varies directly with the extent of unionism, then Fuchs's regression coefficients indicate variations between industries in union wage levels. They would indicate that the wage rates of a completely organized industry would exceed those of a completely unorganized industry by 21 per cent in the "all industry" sample or 25 per cent in the "industrial sector" sample. At the mean values of the unionization variables, this implies average union effects of 7.5 and 12.8 per cent.[4]

The assumptions used in this interpretation are clearly open to question. In particular, there is considerable evidence that union and nonunion firms do not pay the same wages in a given industry. Accordingly, a number of studies have used an alternate interpretation suggested by Lewis. If it is assumed that the excess of union over nonunion wages in a given industry is uncorrelated with the extent of unionism in the industry and also that the level of nonunion wages is uncorrelated with the extent of unionism, then the regression coefficient of the extent of union organization variable can be shown to be an unbiased estimate of log $(1 + \bar{M})$, where $\bar{M}$ is the average union-nonunion wage differential expressed as a proportion of the nonunion wage rate.[5] Accordingly, estimates of the average union-nonunion wage differential are obtained by taking the antilogs of the regression coefficients and subtracting one. This procedure when applied to the Fuchs regression coefficients yields estimates of 24 per cent for the "all industry" sample and 29 per cent for the "industrial sector" sample.

The interpretation suggested by Lewis will be the one used in the following review of the literature. However, two important limitations concerning this interpretation must be kept in mind. First, the estimates may be biased in an unknown direction. For example, it is not clear whether the difference between the union and nonunion wages in a given industry is positively or negatively correlated with the extent of unionism and, accordingly, whether the estimate is biased upwards or downwards.[6]

[3] For example, see Leonard Weiss, "Concentration and Labor Earnings", American Economic Review, Vol. 56, No. 1 (March 1966), pp. 96-117.

[4] The calculation of a true overall union-nonunion differential under these assumptions would require information on the distributions of union and nonunion workers by the extent of union organization in all industries.

[5] See Appendix A.

[6] While there is some reason to believe that the impact of unions on wages, compared to a competitive norm, may be positively associated with the extent of unionism, this does not imply that the difference between the union and nonunion wage rates in an industry will be positively correlated with the extent of unionism.

Second, the parameter being estimated is the actual difference between union and nonunion wages. It is not the difference between union wages and nonunion wages uninfluenced by the presence of unionism. If there are significant spillover effects of unionism on nonunion wage rates, the estimates of the union-nonunion wage differential will understate the actual impact of unionism on wage rates.[7]

With these questions of interpretation being put aside, we can return to a closer examination of Fuchs's results. He found that a nonlinear specification of the extent of unionism variable gave best results. In this specification the value of 0.2 was given to all industries where the extent of unionism was 0.2 or less and the value of 0.6 was given to all industries where the extent of unionism was 0.6 or more. Taking into account the distribution of union and nonunion workers in industries falling into the ranges indicated above, the regression coefficients estimated imply an average union-nonunion differential of approximately 20 per cent.[8]

Ashenfelter and Johnson interpret the Fuchs regression results as indicating a union-nonunion differential of 28 to 35 per cent. However, they do not indicate the basis of their interpretation.[9]

Weiss[10]

Weiss was primarily interested in exploring the relationship between wages and industry concentration. Using data from the one in 1,000 sample of the 1960 U.S. Census of Population, the private annual wage and salary income of individuals was regressed against variables that relate either to their personal characteristics or to the characteristics of the industry in which they work. The industry characteristics included a concentration ratio (CR), the extent of collective bargaining (U), and an interaction term (CR·U). Other industry variables were introduced to control

---

[7] See Appendix A.

[8] See Stanley R. Reber, "The Effects of Unionism Upon Relative Wages" (unpublished dissertation, the University of Chicago, 1970), p. 45.

[9] See Orley Ashenfelter and George F. Johnson, "Unionism, Relative Wages and Labor Quality in U.S. Manufacturing Industries" (Work Paper No. 9, Industrial Relations Section, Princeton University, 1969), p. 2.

[10] Weiss, "Concentration and Labour Earnings", pp. 96-117.

for industry employment growth, industry type (durable and non-durable), establishment size, and proportions of the labour force that are male, skilled, white, in the South, and of non-urban residence. The personal characteristics controlled for through various sets of dummy regressors included race, urban residence, region, size of community, age, years of schooling, family size, status within the family, recent mobility, national origin of parents, hours worked, and the number of weeks worked in the year. The coverage included all persons in specified sex-occupation groups employed for more than 13 weeks in mining, construction, manufacturing, transportation, communications or public utilities. The data was also analyzed excluding the last three "regulated" industries.

With respect to the impact of unionism, Weiss chose to summarize his results by comparing the expected wage for weakly organized industries (U = 50) to the expected wage for highly organized industries (U = 90), assuming mean or arbitrary values for the other independent variables. Using this approach and the sample of unregulated industries, he concluded that unions that organized virtually all their jurisdictions were able to increase earnings by 7 to 15 per cent for craftsmen and 6 to 8 per cent for operatives, compared with poorly organized industries. The spread in the estimates is due to assuming that concentration has either a low or high value. These results and others are summarized in the tabulation appearing on the next page.

When interpreting the Weiss results along the lines suggested by Lewis, a slight modification has to be made because the dependent variable is not in a log form. In this case, the regression coefficients associated with the extent of unionism indicate the average increase in union wages over non-union wages in dollars per year.[11] For the interaction variable $CR \cdot U$, the mean value of CR has been used. Based on this approach, the Weiss regression coefficients for the unregulated industries suggest an average union-nonunion wage differential of 20 to 22 per cent for male operatives and 28 per cent for male craftsmen.[12]

Two additional points concerning the Weiss estimates should be noted. First, there appears to be little relationship between skill level and union effects, at least among production workers. It does appear, however, that the effects of industry unionism upon clerical workers are small. Weiss

[11] It is being assumed that the difference between union and non-union wages and also the level of nonunion wages are both uncorrelated with the extent of unionism. See Appendix A.

[12] Similar calculations for other occupational groups analyzed by Weiss could not be made as the required regression coefficients were not published.

found that once personal characteristics were introduced into the regression equations the relationship between concentration and earnings was rarely significant and as often negative as positive. Moreover, there appeared to be no indication of significant interactions between unionism and concentration.

Percentage Income Advantage of Workers in Industries with High Collective Bargaining Coverage (U = 90) Over Those With Low Collective Bargaining Coverage (U = 50)

OCCUPATION	UNREGULATED INDUSTRIES		ALL INDUSTRIES	
	CR = 20	CR = 60	CR = 20	CR = 60
Craftsmen (male)	8	15	14	9
Operatives (male)	6	8	10	14
Laborers (male)	2	45	13	15
Clerical (female)	–9	6	–8	6

SOURCE: Leonard W. Weiss, "Concentration and Labor Earnings: Reply" The American Economic Review, Vol. LVIII, No. 1 (March, 1968), pp. 181-84.

Throop[13]

Throop estimated the union-nonunion wage differential by regressing the log of industry average earnings against an index of skill mix, an index of urban location, and the extent of union organization in the industry. The industries included in the analysis were bituminous coal mining, contract construction, crude petroleum and natural gas fields, metal mining, quarrying and nonmetallic mining, retail trade, wholesale trade, and two-digit manufacturing industries. The regression coefficients for the extent of unionism variable turned out to be .223 using 1950 data and .260 using 1960 data. Based on the Lewis interpretation, this suggests estimated union-nonunion wage differentials of 25 and 30 per cent respectively.

Rosen (1969a)[14]

Rosen was mainly interested in explaining variations in average weekly hours between industries. A two-equation model was used. In one

[13] A. W. Throop, "The Union-Nonunion Wage Differential and Cost-Push Inflation", The American Economic Review, Vol. LVIII, No. 1 (March 1968), pp. 79-99.

[14] Sherwin Rosen, "On the Interindustry Wage and Salary Structure", Journal of Political Economy, Vol. 77, No. 2, (March/April 1969), pp. 249-73.

equation the supply of hours per week was related to average hourly earnings, other non-wage income, non-pecuniary factors of employment, as well as tastes (region, city size, establishment size, age, race, and marital status). The demand for hours per man was related to factors affecting variable labour costs (average hourly earnings); factors affecting overhead labour costs (age, education, the proportion of non-production labour, marital status, and race); factors affecting short-run adjustments (the rate of change of employment and seasonality); and employment market restrictions (union power as measured by the extent of unionism in the industry). The data on earnings and other variables were standardized according to the industry's occupational distribution. The basic data source was the one in 1,000 sample of the 1960 Census of Population. The sample of industries consisted of 69 observations on two, three, and four-digit industries in mining, construction, manufacturing, transportation, communications and public utilities. From the reduced form of the equations, the regression coefficients of the extent of union organization indicate an average union-nonunion differential of 18 to 25 per cent calculated at the sample mean wage rate.

Rosen (1969b)[15]

This article by Rosen is unique in that it has, as its main focus, variations in union wage effects. In particular, Rosen was interested in:

(i) whether the ability of unions to maintain high wage levels varies with the extent of union organization in an industry, and

(ii) whether indirect or threat effects on nonunion firms vary with the extent of union organization.

The former was dubbed the "wage-coverage" relation, and the latter the "threat-coverage" relation. In the model specified by Rosen, these relations were identified in the ultimate estimating equation.

Average hourly earnings was the dependent variable. The independent variables included a set of three dummy regressors indicating levels of union organization. These dummy variables were also interacted with the level of union organization measured in per cent. Rosen demonstrated that the regression coefficients associated with this rather strange-appearing specification can be used to identify "wage-coverage" and "threat-coverage" relations. Other independent variables included were industry concentration, its interaction with the per cent of the industry organized,

---

[15] Sherwin Rosen, "Trade Union Power, Threat Effects, and the Extent of Organization", The Review of Economic Studies, Vol. XXXVI(2), No. 106, (April 1969), pp. 185-96.

median education, proportion of male workers, proportion of non-white workers, median age, proportion of workers in urban places, proportion of production workers, proportion of skilled crafsmen and kindred workers, proportion of total man-hours worked in the South, and proportion working in establishments of 250 employees or larger. The data were obtained largely from the 1958 U.S. Census of Manufacturers and the 1960 U.S. Census of Population. The observations consisted of 59 two- and three-digit manufacturing industries.

The regression results indicated a strong, positive "wage-coverage" relation. Unions in industries with a coverage in excess of 80 per cent were estimated to have a wage advantage over other unions of from $0.20 to $0.50 per hour, or, in relative terms from 10 to 35 per cent. Contrary to expectations, the "threat-coverage" relation turned out to be negative. That is, nonunion firms in highly organized industries had their wages affected by unionism to a lesser extent than nonunion firms in poorly organized industries.[16]

Earnings showed a positive relation with concentration but a negative relation with the interaction between concentration and the extent of union organization. In Rosen's conceptual framework these results indicated that:

(i) large "threat-coverage" effects may exist in concentrated industries, and

(ii) concentration increases the size of "threat-coverage" effects on nonunion firms to a greater extent than the size of union direct effects.

Rosen used his regression results to estimate an average union-nonunion wage differential comparable to those noted for other studies. The differential turned out to be between 16 and 17 per cent.

Rosen (1970)[17]

In this article, Rosen was primarily interested in exploring the impact of unionism on the occupational wage structure. The estimating equation derived from a carefully specified model had the log of average hourly earnings as the dependent variable. The main unionism variables were the extent of union organization (U) and the product of this variable and the proportion of craftsmen and kindred workers in the industry. In some

[16] Rosen does provide an involved explanation of why this result should not necessarily be considered unusual.

[17] Sherwin Rosen, "Unionism and the Occupational Wage Structure in the United States", International Economic Review, Vol. III, No. 2, (June 1970), pp. 269-86.

equations the square or reciprocal of U was also added as well as the interaction between U and concentration, establishment size, and labour's share of total cost. Other independent variables included urban location, race, age, education, sex, region, and establishment size. The sample consisted of 59 manufacturing industries with data derived from the 1958 and 1963 Census of Manufacturing and the 1/1000 sample of the 1960 Census of Population.

Rosen summarized his main findings as follows:

> If production labor is divided into skilled or not-skilled categories, unionism has widened wage differentials, increasing wage rates of union skilled craftsmen compared with nonunion skilled craftsmen by relatively more than corresponding union-nonunion rates for all other production workers. Further disaggregation of production labor into skilled craftsmen, semi-skilled operatives, and unskilled laborers indicates unionism has probably increased wage rates of unskilled laborers by at least as much and possibly more than that of skilled craftsmen, confirming a result of several other investigations. However, the union-nonunion differential of unskilled labor is significantly higher than that of semi-skilled operatives. The latter effect is quite small though this group constitutes a high proportion of all production workers and the outcome is that unionism has most likely widened the occupational wage structure when all these groups are considered together.[18]

Rosen's analysis also produced other results of interest for this study. In order to capture variations in the union-nonunion differential based on the extent of union organization in an industry, U was entered nonlinearly as U^2 or $1/U$. U^2 was never significant in the log regressions, but often had the expected positive sign. It was often significant and positive in the arithmetic regressions. The reciprocal form was highly significant and had the expected negative sign. In short, the results supported a fairly strong "wage-coverage" relation. Another interesting finding was that effects of the interaction variable ($CR \cdot U$) were generally weakly negative and insignificant.

An overall estimate of the relative effect of unionism on all production workers was obtained by weighting the implied estimates for each skill level group by the proportion of total workers in that group. The estimates range from 25 to 35 per cent and were remarkably consistent from equation to equation. However, Rosen believed these estimates to be subject to an upward bias due to the positive correlation between the

---

[18] Ibid., p. 269.

extent of union organization and the union-nonunion differential. Using a specification that allowed the differential to vary produced an average differential of about 17 to 22 per cent at the mean level of U.

Raimon and Stoikov[19]

The authors in this article were primarily interested in examining wage spillover effects from production to nonproduction workers. Their model related average weekly earnings for either production or nonproduction workers to an index of industry skill level, the extent of union organization (U), a concentration ratio (CR), an interaction variable (U·CR), the median years of schooling adjusted for skill mix, and the proportion of employees in large establishments who are Negro, age 24 or under, in the South, and in rural areas. Two samples of various groupings of three- and four-digit industries were used. One sample consisted of industries in mining, manufacturing, construction, transportation, communications, and public utilities. The second sample was confined to manufacturing. The data were extracted, in the main, from BLS publications covering the year 1960.

The implied union-nonunion wage differential for production workers, evaluated at an assumed mean value of .40 for the concentration ratio, turns out to be 25 per cent for the large sample and 31 per cent for the small sample. For nonproduction workers, the authors concluded that the regression coefficients indicating union impact were not significantly different from zero.

For production workers, the coefficient attached to the concentration variable (CR) was invariably positive, but the interaction term (U·CR) was consistently negative. In the equations covering nonproduction workers, both these variables were not statistically significant.

Hamermesh[20]

Hamermesh specified a model in which union effects for both white- and blue-collar workers can be determined. His model was derived from relative supply and demand conditions for both these types of labour. The reduced form equation had the log of the ratio of white- and

---

[19] Robert L. Raimon and Vladimir Stoikov, "The Effect of Blue-Collar Unionism on White-Collar Earnings", Industrial and Labor Relations Review, Vol. 22, No. 3 (April 1969), pp. 358-74.

[20] Daniel S. Hamermesh, "White-Collar Unions, Blue-Collar Unions and Wages in Manufacturing", Industrial and Labour Relations Review, Vol. 24, No. 2 (January 1971), pp. 159-70.

blue-collar wages being related to a measure of clerical work "intensity", education levels, region, the extent of union organization among blue-collar workers, and the extent of union organization among white-collar workers. In this formulation, the regression coefficients of the last two variables provide estimates of the white- and blue-collar union-nonunion wage differentials. The model was estimated using manufacturing wage rate data for metropolitan areas collected in the BLS annual wage rate survey in the years 1960-61, 1963-64, and 1966-67. The equations were run separately for a number of pairwise (white- and blue-collar) groups of sex-occupational classifications and also on a pooled basis. The results indicated a union-nonunion differential of 15 per cent for male blue-collar occupations and 21 per cent for females. However, for white-collar workers, the differential was quite small, roughly 5 per cent.

Pencavel[21]

Although Pencavel was primarily interested in explaining the relations between wages, specific training, and labour turnover, his model does provide evidence on the impact of unions. In one equation, the log of industry average wages was related to the extent of union organization, the ratio of males to females, the quit rate, and the median years of education. The sample consisted of nineteen two-digit industries for the years 1959 and 1960. The average union-nonunion wage differential derived from this equation was 27 per cent.

Rapping[22]

Rapping specified a model in which the log of the industry average wage of male labourers was the dependent variable. Independent variables included race, location, the rate of growth of employment, and the extent of union organization. Ability-to-pay was captured by two alternate measures of monopoly rents per man-hour. The first specification used profits per man-hour and assets per man-hour as independent variables. The second used concentration times value added per man-hour. To determine whether unions share in monopoly profits, all these measures of ability-to-pay were also interacted with the extent of union organization in

[21] John H. Pencavel, "Wages, Specific Training and Labor Turnover in U.S. Manufacturing Industries", (Institute for Mathematical Studies in the Social Sciences, Stanford University, June 1971, mimeographed).

[22] Leonard A. Rapping, "Monopoly Rents, Wage Rates and Union Wage Effectiveness", The Quarterly Review of Economics and Business, Vol. 7, No. 1 (Spring 1967), pp. 31-47.

the industry. The data consisted of average values for 19 two-digit manufacturing industries in 1959.

The results suggest that there is a positive rents per man-hour effect on wage rates independent of unionism. This finding was supported by both the profit and concentration measures of rents per man-hour. There was also limited support for the hypothesis that collective bargaining was more effective in monopolistic industries. However, the effectiveness of unionism was mainly related to factors other than the presence of product market monopoly.

The implied average union-nonunion wage differential ranged from 19 to 66 per cent depending on the exact specification of the model.

Studies Using Data on Individual Workers or Establishments

Stafford[23]

With the 1966 Survey of Consumer Finances conducted by the Survey Research Center of the University of Michigan, Stafford was able to utilize the union status of individuals as a basic "union" variable in explaining wages. After certain exclusions, the survey yielded 1,111 observations. The family head's annual earnings was the dependent variable. The independent variables were sets of dummy regressors that indicated the union status of the individual, weeks worked, schooling, age, occupation, hours worked per week, city size, region, race, supervisory status, and industry. The last set of dummy variables divided industry by major divisions (e.g. manufacturing, trade, services, etc.).

When all occupations were considered together, the union-nonunion differential was 16 per cent. However, based on the more appropriate procedure of running the regressions separately for each occupation the results were as follows:

Union-Nonunion Differentials in Per Cent for Four Occupational Groups

Occupation	No. in Sample	The Differential
Operatives	281	26
Craftsmen	324	24
Labourers	165	52
Clerical & Sales	126	18

[23] F. P. Stafford, "Concentration and Labor Earnings: Comment" American Economic Review, Vol. LVIII, No. 1 (March, 1968), pp. 174-81.

Johnson and Youmans[24]

The authors were primarily interested in exploring union effects by age and education. In their estimating equation the dependent variable was the log of an individual's wage rate. The independent variables included the union status of the individual, his age, and education, plus a set of variables allowing interaction between these variables and unionism. Other variables controlled for race, urban location, and region. The data used were similar to that of Stafford's. They were taken from the 1965 and 1966 Survey of Consumer Finances and included all male heads of households in blue-collar occupations (craftsmen, operatives and labourers). The sample included 1,950 individuals.

The results indicated a union-nonunion wage differential of 34 per cent. There was also evidence that the union impact was greater for less educated workers.

Oaxaca[25]

In this paper the author was concerned with measuring the extent of sex discrimination. However, the union status of the individual was also one of the factors controlled for in the wage determination equation. The model involved regressing the log of the hourly wage rate of the individual against sets of dummy variables indicating work experience, education, industry, occupation, migration, marital status, size of urban area, and region. The "industry" regressors were based on "major divisions." The data source was the 1967 Survey of Economic Opportunity. About 60,000 individuals in 30,000 households were covered.

The estimated union-nonunion wage differentials were as follows: white males – 11 per cent, white females – 15 per cent, black females – seven per cent, and black males – 21 per cent.

Ashenfelter[26]

This article by Ashenfelter presented evidence on the impact of unionism on white-black wage differentials. He also used the 1967 Survey of Economic Opportunity and variables that were identical to Oaxaca's.

---

[24] George E. Johnson and Kenwood C. Youmans, "Union Relation Wage Effects by Age and Education", Industrial and Labor Relations Review, Vol. 24, No. 2 (January, 1971), pp. 171-79.

[25] Ronald L. Oaxaca, "Male-Female Wage Differentials in Urban Labor Markets", (Working Paper No. 23, Industrial Relations Section, Princeton University, n.d.).

[26] Orley Ashenfelter, "Racial Discrimination and Trade Unionism", The Journal of Political Economy, Vol. 80, No. 3, Part I (May/June, 1972), pp. 435-64.

However, in Ashenfelter's specification the union status of the individual was interacted with dummy variables indicating various occupational-industry categories. The results were as follows:

Union-Nonunion Wage Differentials in Per Cent by Occupation, Race, and Industry

Occupation	White Workers		Black Workers	
	Non-Construction	Construction	Non-Construction	Construction
Craftsmen	3	40	13	52
Operatives	15	44	22	33
Labourers	19	48	32	46

For white workers in white-collar jobs the differentials were negligible except for professional workers (12 per cent). For black workers in white-collar jobs the differential ranged from 12 to 32 per cent.

Ashenfelter also provided a more detailed breakdown of the differential by industry for white blue-collar workers. This is given in the following tabulation:

Union-Nonunion Wage Differentials in Per Cent by Occupations and Industry for White Workers

Industry	Occupation		
	Craftsmen	Operatives	Labourers
Construction	40	44	48
Durable Manufacturing	-2	12	15
Non-Durable Manufacturing	4	13	12
Transportation, Communications, Utilities	2	14	22
Other Industries	17	34	30

Ashenfelter reports the overall union-nonunion wage differentials to be 10 per cent for white males and 21 per cent for black males. This estimate was obtained by using weights for each of the occupation-industry categories. For females, the comparable differentials are reported as 15 and 7 per cent for whites and blacks respectively.

Bailey, King and Schwenk[27]

This study is unique in that the establishment was the unit of observation. The dependent variable was straight-time total compensation per man-hour worked for production employees. The independent variables reffered to characteristics of the establishment or its industry (generally two-digit). The establishment variables included union status, employment size, and location. The industry variables included concentration, skill mix, a capital labour ratio, and value added per employee. The sample was confined to the manufacturing sector and consisted of 1,149 observations for the year 1968.

In what the authors judged to be their best equation, the union status coefficient indicated a differential of 50 cents per hour. At the mean value for the dependent variable, this indicated a differential of 15 per cent. The authors found no relation between concentration and wages. Nor did they find any union interaction effects. This result was reported as follows:

> We were particularly interested in whether interaction among unionization, concentration, and size might have an influence on wage rates. However, in the several tests that were run, we found no evidence that interaction among these factors should have any effect on the analysis. An additional interaction test was performed by introducing an industry unionization variable and testing its interaction with establishment unionization. No significant result was obtained.[28]

Reber[29]

This study differs from the others discussed in this chapter in that the author was primarily interested in developing estimates of the ratio of union to nonunion wages for individual industries. To do this, Reber examined 101 reports of the BLS covering the period 1945 to 1968. Each of the reports provided information on average hourly earnings for various occupations in a given two- to five-digit industry as of a given date. The industries were primarily from manufacturing. These average hourly earnings were also classified by region, the union status of the establish-

---

[27] William R. Bailey, Harlan W. King and Albert E. Schwenk, "Wage Differentials and Establishment Size, Union Status and Concentration", American Statistical Association, 1970 Proceedings of the Business and Economic Statistics Section, (Washington, D.C., 1971), pp. 393-98.

[28] Ibid., p. 397.

[29] Stanley R. Reber, "The Effects of Unionism Upon Relative Wages."

ment, and, in some cases, establishment and community size. By applying appropriate weights to the ratios of union to nonunion wages in each of the cells, an overall ratio of union to nonunion wages in a given industry as of a given date was calculated. Disregarding dates, the average industry differential was estimated to be from 10 to 15 per cent.

Reber also analyzed the differentials for particular dates and industries to determine if they exhibit any patterns. In contrast to Lewis, he found it impossible to identify any strong relationship between date and the size of the union-nonunion wage differentials in the period 1945 to 1968. This result was interpreted as being in conflict with the generally held view that, due to the rigidity of union wages, there will be a tendency for the differential to vary according to actual and expected rates of inflation.

Inter-industry variations in the differential were found to be substantial. When possible causes of these variations were examined, it was found that the differential increased significantly with the extent of union organization. The effect was estimated to rise some 16 to 27 per cent as the extent of union organization increased from 0 to 100 per cent. The relationship appeared to be nonlinear with the main increase coming in the unionization range somewhat above 50 per cent. No relationship was found between the size of the differential and industry concentration.

An examination was also made of the impact of unionism on the occupational wage structure. Unionism appeared to cause a narrowing of occupational wage differentials but the relations were not statistically significant.

Summary and Assessment

The studies reviewed in this chapter are not strictly comparable. There are variations in methodologies, concepts, and the extent to which factors other than unionism can be controlled for in the wage determination equations. Important differences also exist in the size and quality of the samples and the time period covered. However, despite these differences, some general statements about the results can be made.

The global estimates of the union-nonunion wage differential demonstrate wide diversity. These estimates for production workers, based on the interpretation suggested by Lewis, are displayed in Table 1. In the case of the studies based on industry aggregated data, the estimates, in the main, fall in the range of 20 to 30 per cent. This is significant as it is double the range suggested by Lewis. The difference appears to be too large to be accounted for by differences in reference periods for the studies.

Table 1

Global Estimates of Union-Nonunion Wage Differentials for Production Workers in the United States

Author	Coverage	Union-Nonunion Wage Differential in Per Cent
Industry Aggregated Data		
Fuchs (1968)	1960, non-agricultural industries	20
Weiss (1966)	1960, mining, construction, manufacturing, transportation, communications and public utilities	Male Operatives: 20-22
Throop (1968)	1960 & 1950, mining, manufacturing, contract construction and trade	1950: 25 1960: 30
Rosen (1969) (1)	1960, mining, construction, manufacturing, transportation, communications and public utilities	18 – 25
Rosen (1969) (2)	1958, manufacturing	16 – 17
Rosen (1971)	1960, manufacturing	17 – 22
Raimon & Stoikov (1969)	1960, mining, manufacturing, construction, transportation, communications and public utilities	25 – 31
Hamermesh (1971)	1960-61, 1963-64, and 1966-67, manufacturing	15 – 21
Pencavel (1971)	1960, manufacturing	27
Rapping (1967)	1959, manufacturing	Male labourers: 19-66
Disaggregated Data		
Stafford (1968)	1966, Survey of Consumer Finances: male family heads all non-agricultural industries	Craftsmen: 24 Operatives: 26 Labourers: 52
Johnson & Youmans (1971)	1965 and 1966, Survey of Consumer Finances: male family heads in blue-collar occupations all non-agricultural industries	34
Oaxaca (1971)	1967, Survey of Economic Opportunity all industries in urban areas	White Males: 11 White Females: 15 Black Males: 21 Black Females: 7
Ashenfelter (1971)	(As above)	White Males: 10 Black Males: 21
Bailey, King & Schwenk (1968)	1968, BLS establishment survey in manufacturing	15
Reber (1970)	1945-1968, BLS industry surveys (mainly manufacturing).	10 – 15

One possible explanation of the difference is that the studies using industry aggregated data are subject to an upward bias. As mentioned above, the interpretation suggested by Lewis only gives an unbiased estimate when there is no correlation between extent of union organization by industry and either the size of the differential by industry or the size of threat effects by industry. As there was some evidence of a positive "wage-coverage" relation in two of the studies done by Rosen and also the study by Reber, it would appear that the indicated estimates are biased upwards.

However, it is uncertain whether this provides a total explanation. Not enough information is known to calculate the possible size of this bias. Moreoever, the study by Bailey, King and Schwenk found no relation between the size of the differential and the extent of union organization. Also, Rosen's finding that threat effects on nonunion firms are negatively related to the extent of union organization indicates a source of downward bias in the estimates. In short, there is some doubt concerning the confidence that can be placed in the studies using industry aggregated data.

This difficulty suggests that more weight should be placed on the studies using the individual or the establishment as the unit of observation, as they are not subject to the same problems of interpretation. But the estimates derived from these studies display considerable diversity. The two studies based on the Survey of Consumer Finances show quite large estimates while the two studies based on the 1967 Survey of Economic Opportunity and the study of Bailey *et al* show relatively low estimates.

From all this, one might consider that recent evidence suggests global estimates of the union-nonunion wage differential for production workers that are somewhat higher than those proposed by Lewis. However, it would be hard to make a "best estimate."

The studies do shed some light on variations in union impact, but the findings are not always consistent. Reference has already been made to the support for a positive "wage-coverage" relation. Unions that are able to organize a large part of their jurisdictions apparently are relatively more successful than other unions. However, the evidence for such a relationship is far from complete. The two studies using the ideal procedure of controlling for both the union status of the establishment and the extent of industry unionism, came to opposite conclusions. Only the study of Rosen has explored the relationship between unionism and nonunion wages. The negative "threat-coverage" relation indicated by this study must be viewed as only a tentative result.

The evidence on the impact of concentration is mixed. Rosen, Raimon

and Stoikov, and Rapping all reported positive influences of concentration on earnings. Weiss also reported an initial positive relationship but this disappeared when personal characteristics were introduced into the wage determination equation. The study of Bailey *et al,* using disaggregated data, found no relationship between concentration and earnings. There is agreement between the studies on at least one point. No evidence of a significant interaction between union impact and concentration was found in any of the studies. If anything, the results suggest that the interaction may be negative.

The available studies tend to support the hypothesis that unionism narrows occupational wage differentials, at least for production workers. This is evident in the studies using disaggregated data by Ashenfelter, Stafford, and Reber. This hypothesis was also supported indirectly by Johnson and Youmans who found that unions benefit less educated workers to a greater extent than more educated workers. The verdict, however, was not unanimous. Weiss's results show no relation between skill level and the size of union-nonunion differentials and Rosen, using a measurement methodology based on highly aggregated data, concluded that unionism actually widens the occupational wage structure.

The three studies that report union-nonunion wage differentials by sex all find that the female differential is larger. The female advantage, however, is not large being consistently in the range of 4 to 6 per cent in each study.

Regarding white-collar workers, the studies are all in agreement. Both unionism among white-collar workers and unionism among associated blue-collar workers appear to have a small impact on white-collar wages.

One final characteristic concerning all these studies merits emphasis. They all aim at measuring the observed difference between union and nonunion wages. None of the studies attempts to measure the difference between actual union wages and what these wages would be in the absence of unionism from the industry or economy as a whole.

CHAPTER II

THE DATA AND THEORETICAL FRAMEWORK

As has been shown, the existing knowledge of union-nonunion wage differentials is far from complete. First no studies exist concerning the Canadian experience. Even for the United States the available estimates of the general level of these differentials display considerable diversity. A number of recent studies suggest differentials that are considerably in excess of the prevailing benchmark estimate of 10 to 15 per cent. However, some of the estimates and, indeed, to some extent the benchmark itself, are open to question due to their reliance on highly aggregated data. Moreover, there are only a few quantitative estimates of variations in union impact. A particularly glaring gap in the existing research is that the actual size of union spillover effects on the wages of nonunion workers has remained virtually unexplored.

The purpose of this study is to examine the Canadian experience regarding these questions with the aid of a new data set and analytical approach. In particular, the study will analyze:

(i) the mean union-nonunion wage differential for production workers,

(ii) sources of variation in the differential related to industry characteristics,

(iii) variations in the differential by sex and skill levels, and

(iv) spillover effects of unionism on nonunion white- and blue-collar workers.

In this chapter, a model of wage determination, appropriate for considering these topics, is developed.

The Data Set

The analytical approach selected for this study has been considerably influenced by the unique data set that is available. Therefore, it is useful to describe it in broad terms at the outset.

A tape, available from the 1969 Canada Department of Labour Occupational Wage Rate Survey, provides information on union status, detailed occupational wage rates, employment, location, industry, and other information for individual establishments in Ontario. Almost all establishments with twenty or more employees are included in the survey. Although the survey includes a wide range of industries, only those within manufacturing are considered in this study. Non-manufacturing industries

had to be excluded due to the unavailability of supplementary data required for the analysis. The 1969 survey returns were selected for analysis because they were the latest available when this study was initiated and resource limitations made it impractical to consider a number of years. Moreover, there have been a number of changes in the form of the question dealing with the union status of establishments, thereby complicating a mixed cross-section and time series approach.

The survey solicits wage rate and employment information for a set of specified occupations. Although most of these are industry specific, and hence difficult to analyze, some, including various office, service, and maintenance jobs, are found in a wide range of industries. For given broad skill level categories, those inter-industry occupations for which the establishment response was greatest were chosen for this study. The questionnaire also asks for a basic wage rate, that is, the lowest hourly wage rate after a probationary period paid to labourers or equivalent unskilled employees. As virtually all establishments report such a rate, this "occupation" is included in the analysis.

The survey attempts to keep homogeneity within the narrowly defined occupational classes. For each occupation, the questionnaire contains a short description of work characteristically performed. Separate information for male and female employees is requested. Also the respondents are instructed to exclude learners, apprentices, beginners, trainees, helpers, and probationary, part-time, supervisory (including foremen and lead hands) and temporary employees and helpers. In the reporting of wage rates, retroactive wage adjustments, overtime premium payments, and shift differentials are excluded. Actual wage rates paid to various employees are given for each of the specified occupations. The basic rates are essentially hiring rates for entry level jobs.

Ideally one would want to know whether the occupations selected for analysis are covered by a collective agreement in a particular establishment. Unfortunately, the data on unionism provided by the survey are not quite that precise. The questionnaire asks: "Do you have any written collective agreements (in force or being negotiated) with one or more organizations representing your employees?" Separate provision is made for answering the question for non-office and office workers. An establishment has been classified as "union" for a given blue-collar occupation if the answer for non-office employees is "yes." A similar approach was used for the white-collar jobs.

This procedure undoubtedly contains some error as it is conceivable that a small craft unit of non-office employees could be represented by a union while the main body of production workers could be unrepresented.

In these situations, the establishment would be classified as "union" even though the occupation in question was not covered by a collective agreement. However, given the pattern of union organization in the province, this source of error is probably not important. Outside construction, almost all employees are organized on an "all production employee" basis rather than the narrow craft unit basis. Even in those establishments where craft units exist, it is only in exceptional cases that the associated production employees are not organized.

Basic data on employment, wages, and union status for the selected inter-industry occupations and basic rates are provided in Table 2. It can be seen that, particularly for the blue-collar occupations, there are a large number of reporting establishments in each of the cells. Only in the case of unionized white-collar occupations is there found fewer than 100 observations. For this reason, the survey can be viewed as a potentially powerful data base for considering union-nonunion wage differentials.

The usefulness of the survey is also enhanced by its high level of coverage. The Census of Manufacturers reports 5,783 establishments in Ontario with fifteen or more employees. These accounted for 547,667 production employees.[1] As can be seen from Table 2, the occupational wage rate survey, although limited to establishments with twenty or more employees, provides data on 3,748 establishments and 507,665 non-office employees. For the larger establishments, the survey is virtually complete. Although there are many establishments with fewer than fifteen employees, they account for only 4 per cent of total production worker employment.

In the model to be developed in this chapter, an attempt is made to explain the inter-establishment wage rate variation reported by the survey within each of the selected occupations and basic rates. The explanation is to be in terms of quantifiable establishment and industry characteristics, including unionism.

Union-Nonunion Wage Differentials – The Concept

The concept of a union-nonunion wage differential is not intuitively obvious, and, in fact, a number of definitions are possible and are used in the literature. The simplest is the extent to which the wages of union workers exceed the wages of nonunion workers.[2] The wage rates

[1] Dominion Bureau of Statistics, Manufacturing Industries of Canada, Type of Organization and Size of Establishments, 1967, Catalogue No. 31-210 (Ottawa: Information Canada, 1972).

[2] See, for example, Sandra L. Mason, "Comparing Union and Nonunion Wages in Manufacturing," Monthly Labour Review (May 1971), pp. 25-26.

TABLE 2
UNION AND NONUNION WAGE RATES IN MANUFACTURING INDUSTRIES IN ONTARIO 1969

Occupation	Union			Nonunion		
	Establishments	Employees	Average Wage Rate	Establishments	Employees	Average Wage Rate
Basic Rates (a)			$ per hour			$ per hour
Male	1,950	395,705(b)	2.54	1,798	111,960(b)	2.10
Female	1,950	395,705(b)	2.18	1,798	111,960(b)	1.78
Unskilled Blue-Collar						
Cleaner	970	4,477	2.70	547	1,054	2.29
Labourer, Production	779	12,086	2.57	446	2,795	2.16
Labourer, Non-Production	510	5,151	2.72	225	891	2.32
Semi-Skilled Blue-Collar						
Shipper	1,337	6,800	2.77	1,064	2,700	2.51
Industrial Truck Operator	822	6,261	2.98	325	837	2.69
Truck Driver, Heavy Truck	483	2,761	3.20	339	1,144	2.73
Skilled Blue-Collar						
Electrician	708	3,829	3.78	187	526	3.71
Machinist	518	3,089	3.61	220	645	3.42
Mechanic (Machine Repair)	816	5,686	3.55	415	1,092	3.09
Welder	446	2,414	3.53	182	525	3.17
White-Collar Female			$ per week			$ per week
Bookkeeping Machine Operator	45	91	92.83	765	1,023	86.07
Junior Typist	84	455	84.58	913	2,024	74.53
Senior Secretary	125	385	122.02	1,358	3,053	107.83
White-Collar Male						
Senior Accounting Clerk	88	179	135.10	696	1,299	136.86
Order Clerk	70	192	123.94	817	1,725	121.06
Senior Clerk	85	421	137.73	589	2,106	147.04

(a) The basic rate is defined as the rate paid to labourers or equivalent unskilled employees after termination of a learning or probationary period, if any.

(b) Relates to total non-office employment in the establishment.

SOURCE: Canada Department of Labour, Occupational Wage Rate Survey.

embodied in Table 2 provide such a comparison. The studies using this definition are primarily interested in description rather than analysis. They can document the relative position of union workers, but no inference concerning the impact of union status can be made as it is well recognized that other factors can have an influence on wage levels.

For the purposes of analysis, the most widely used definition of the union-nonunion differential is the extent to which the average wage of union workers exceeds the average wage of comparable groups of nonunion workers.[3] The key word in this definition is comparable. This requires that any data being used must be adjusted for factors other than unionism that may affect wages before the size of the differential can be determined. In studies that embrace a number of industries, the adjustments are typically carried out within a multiple regression framework.

What factors other than unionism should be taken into account? The universal practice in this regard is to introduce variables that competitive labour market theory suggests as influences on wage rate levels. In effect, comparisons are being made between groups of workers that according to the competitive model, should receive the same wage rates.[4]

Even though the "comparable group" definition is widely used, it suffers from a major defect when taken as a measure of union impact, as it fails to recognize that the wages of nonunion workers can be affected by the presence of unionism. A more comprehensive definition of the differential would be the extent to which the wages of a group of union labour exceed the wages of a comparable group of labour where unionism has no impact on wages. This definition, while being analytically interesting, borders on being non-operational. The problem arises because all wages in the economy are at least indirectly affected by the presence of unionism. Compared to the competitive norm, if unionism has a positive wage impact, union produced goods and services will become relatively expensive, and hence, there will be a shift in demand towards the products of nonunion firms. Also, labour displaced by the relatively high union wages will increase the labour supply to other parts of the economy. There may even be a change in the relative price and demand for capital. In

[3] This definition is implicit in virtually all the studies reviewed in Chapter 1. It is explicitly made by Albert Rees, The Economics of Trade Unions, (Chicago: The University of Chicago Press, 1962), p. 73.

[4] It might be argued that this procedure is not appropriate as, in the absence of unionism, the labour market would still be imperfectly competitive. See Robert M. MacDonald, "An Evaluation of The Economic Analysis of Unionism," Industrial and Labour Relations Review, Vol. 19, No. 3 (April 1966), pp. 335-47. This point is discussed below.

short, unionism is likely to alter supply and demand conditions in all markets in the economy. Accordingly, the definition of the union-nonunion wage differential given in this paragraph appears to be in some trouble due to the inability to directly observe nonunion wages that are totally uninfluenced by unionism.

This difficulty can be bypassed if a distinction is made between the primary and secondary effects of unionism. Secondary effects are those wage changes due to induce shifts in the supply and demand for labour throughout the rest of the economy produced by the establishment of a positive union wage differential in a given industry. Primary effects are all the wage effects in a given industry associated with the presence of unionism in that industry. These primary effects include the spillover effects of unionism in a given industry on the wages paid by the nonunion firms in the industry. As secondary unionism effects are not directly observable or easily estimated, it is convenient to couch a definition of the union-nonunion wage differential in terms of primary effects only. Accordingly, it can be defined as the extent to which the wages of organized workers exceed those paid to comparable unorganized workers where unionism has not had a primary effect on wages, that is, where spillover effects are nil.

The distinction between the primary and secondary effects of unionism helps clarify the sense in which the union-nonunion wage differential as defined immediately above can be interpreted as a measure of the union impact on wages. If a small scale change in the level of unionization is being considered, say a change in the union status of an establishment or the degree of unionization in a single industry, the level of unionization elsewhere being held constant, the induced secondary effects averaged across all other industries are likely to be very small and, most likely, can safely be ignored. However, if large scale changes in the level of unionization are being considered, say the effect of unionization throughout the economy as a whole, the secondary effects cannot be ignored. In this sense, this definition of the union-nonunion differential only provides a limited measure of union impact.

Put in other words, the union-nonunion wage differential that has been defined provides a partial equilibrium measure of the impact of unionism and does not indicate the impact of all unionization throughout the economy on the absolute wages of either union or nonunion workers. To attempt to measure the total impact of all unionism would require the consideration of a complex set of of general equilibrium adjustments. In effect, an estimate would have to be made of the market outcome of a reallocation of resources from the nonunion to the union sectors of the

economy, along with the effect of a change in the relative price of capital and labour. This is beyond the scope of this study.[5]

In summary, the calculation of union-nonunion wage differentials involves more than a simple comparison of the wages of organized and unorganized workers. As a minimum, the differential must be defined as the difference in wage rates between comparable groups of union and nonunion labour. Although somewhat misleading this will be called the observed union-nonunion wage differential. A more satisfactory definition, for the purpose of analyzing union impact, is the difference between the wages of organized workers and those of comparable unorganized workers where unionism has not had a primary or spillover effect on wages. This will be called the actual union-nonunion differential. The distinction between these two definitions is clarified in the development of the model below.

The Model

In this analysis, the variable to be explained is the wage level of the establishment. There are two kinds of establishments; those that have their wages determined through collective bargaining and those that do not. The wage level in a union establishment, W_u, and the wage level in a nonunion establishment, W_n, can be viewed as comprised of two elements, as follows:

$$W_u = W_c + \alpha_u \text{ , and} \qquad (1)$$

$$W_n = W_c + \alpha_n \qquad (2)$$

where W_c is what the wage would be in the absence of any primary impact of unionism, but with the general level of unionism in the economy held constant; α_u is the primary effect of unionism on the union establishment, and α_n is the primary impact of unionism on a nonunion establishment. For the purposes of estimation, W_c is represented by $\beta X+e$ where β is a vector of fixed coefficients, X is a vector of variables apart from unionism that affect wages and e is an independent normally distributed random error term with a constant variance and a zero mean. The set βX includes a

[5]For a highly aggregated approach to this measurement problem, see Harry G. Johnson and Peter Mieszkowski, "The Effects of Unionization on the Distribution of Income: A General Equilibrium Approach", The Quarterly Journal of Economics, Vol. LXXXIV, No. 4 (Nov. 1970), pp. 539-61.

constant term. As $\beta X+e$ can be assumed to be the same for both union and nonunion establishments, (1) and (2) can be combined to give:

$$W = \beta X + \alpha_u CA + \alpha_n(1-CA) + e \qquad (3)$$

where W is the wage level in any establishment and CA is a dummy variable having a value of one if the establishment is covered by a collective agreement and zero otherwise.

In order to adapt (3) for empirical use, some assumptions must be made about the nature of α_u and α_n. The simplest assumption is that both are constant, that is:

$$\alpha_u = \alpha_{u_0} \quad \text{and} \qquad (4)$$

$$\alpha_n = \alpha_{n_0} \qquad (5)$$

Substituting into equation (3) gives

$$W = \beta X + \alpha_{u_0} CA + \alpha_{n_0}(1-CA) + e \qquad (6)$$

The model in this form can not be estimated using standard regression techniques as the two dummy variables CA and (1-CA) and the constant term of βX are not linearly independent. However, with a slight rearrangement the model can be put into a form that permits estimation, namely:

$$W = \alpha_{n_0} + \beta X + (\alpha_{u_0} - \alpha_{n_0})\, CA + e \quad (7)$$

In a regression equation corresponding to (7), the coefficient attached to the regressor CA would be interpreted as the observed union-nonunion wage differential, that is, the difference between the wage levels in union and nonunion establishments after controlling for other sources of wage variation. Regression equations conforming to this formulation will be labelled as Variant I.

The assumption that α_u is constant is highly restrictive. In fact, exploring variations in α_u is one of the principal aims of this study. Accordingly, it can be assumed that $\alpha_u = \alpha_{u_0} + \alpha_{u_i} Y_i$ where α_{u_i} is a vector of fixed coefficients and Y_i is a vector of variables that are determinants of union effectiveness. Making the appropriate substitutions leads to a model of the following general form:

$$W = \alpha_{n_0} + \beta X + (\alpha_{u_0} - \alpha_{n_0})CA + \alpha_{u_i} Y_i\, CA + e. \qquad (8)$$

Regression equations corresponding to (8) will be labelled Variant II. Again the coefficients must be interpreted as observed union nonunion wage differentials. In effect union workers are being compared with comparable groups of nonunion workers who, however, may be subject to union primary effects.

The model can be extended one step further by dropping the assumption that α_n is constant. If instead, it is assumed that $\alpha_n = \alpha_{n_0} + \alpha_{n_j} Z_j$, the model becomes

$$W = \alpha_{n_0} + \beta X + (\alpha_{u_0} - \alpha_{n_0}) CA + \alpha_{u_i} Y_i \cdot CA + \alpha_{n_j} Z_j (1 - CA) + e. \quad (9)$$

This formulation will be called Variant III.

The model in this form can be estimated as long as the variables that are a part of $\alpha_{u_i} Y_i \, CA$ and $\alpha_{n_j} Z_j (1 - CA)$ are linearly independent. If this condition is satisfied and it is possible to specify conditions under which α_n is likely to be zero or close to zero, the model can be used to define what may be called the actual union-nonunion wage differential. This can be done by comparing the expected wage rate for a nonunion establishment under conditions where the spillover influence of unionism on nonunion wage rates is likely to be negligible to the expected wage for comparable union establishments.

This, in essence, is the strategy that will be followed. In the rest of this chapter, the available theory is reviewed in order to specify what must be included in vectors X (control variables), Y_i (union impact variables) and Z_j (spillover effect variables). There is also some discussion of functional form and measurement problems.

Control Variables

If, within a particular labour market, it is assumed that:

- (i) labour within a particular occupational group is homogeneous from the employer's point of view;
- (ii) non-wage conditions of employment (pecuniary or otherwise) do not vary significantly between employers;
- (iii) the costs of information and labour market employee adjustments are zero; and
- (iv) there are no institutional restraints on the clearing of markets;

then the competitive hypothesis, based on the assumption of profit and utility maximization, predicts that a single wage rate will prevail. In particular, there would be no reason for wage rates in a given occupation to vary systematically among industries within a local labour market.

If the simplifying assumptions are dropped, then the predictions of

competitive theory are modified. No longer would one expect a single wage rate to prevail, but rather the prediction is that there would be a range of wage rates for a given occupational class reflecting compensating or "equalizing" differentials among jobs and people, and short-run wage adjustments to supply and demand shifts.

Location

The labour supply conditions for particular occupations can be expected to vary geographically due to short-run barriers to inter-regional mobility and perhaps locational preference. Similarly, within a given region there may be a systematic difference in wage rate levels between urban and rural employment reflecting such factors as locational preference, differences in costs of living, and perhaps incomplete market adjustments to varying growth rates in labour supply and demand. In order to capture these geographic labour supply and demand influences, two variables are introduced: R, a measure of regional location, and C, a measure of city location. R is defined by a set of dummy regressors that effectively divide the province into five regions, and C is a dummy variable having a value of one if the establishment is located in a major metropolitan area (Hamilton, London, Ottawa, Toronto or Windsor), zero otherwise.

Plant Size

There is some precedent in the literature for using a "size of establishment" variable as a proxy for certain "compensating" differentials. The argument is that large plants need more formal rules and workers who are willing to tolerate "regimentation". Also, because of interdependence in large production units and the fact that erratic performance or excessive absenteeism may be especially costly for large plants, there exists an incentive for large plants to set higher standards for dependability than smaller ones. Similarly, the cost impact of temporary labour shortages might be proportionately greater for larger plants.[6] Another possible reason for a positive relationship between wages and plant size has been suggested by Stigler. In a world where information is not costless, wage rates and search activities are substitutes for the employer. "The more efficiently he detects workers of superior quality the less he need pay for such quality."[7] The employer in a small plant has

[6] For a more detailed discussion of these a priori arguments, see Stanley H. Masters, "An Interindustry Analysis of Wages and Plant Size", The Review of Economics and Statistics, Vol. LI, No. 3 (Aug. 1969), pp. 341-45.

[7] George J. Stigler, "Information in the Labour Market," Journal of Political Economy (Supplement, October 1962), p. 101.

an advantage in that he can closely observe the performance of his employees whereas the large employer must rely on expensive and indirect rating practices. As a result the small employer has a cost advantage in judging quality and need not pay wages as high as those in large plants.

A positive relationship between wages and plant size might also arise due to the omission of certain labour quality variables from the analysis. Much to their surprise, Shultz and Rees[8] found no significant relation between establishment size and wage levels within occupations after introducing a number of labour quality variables into their analysis. They remark:

> Since there is in general a negative relation between turnover and establishment size, one might expect larger establishments to have employees with greater average seniority, and hence higher wages for a given wage structure by length of service . . . [T]he proposition stated about the relation between seniority and establishment size is supported by [. . . the simple correlation coefficients between the natural logarithm of seniority and the natural logarithm of establishment size for each occupation]; nine of the twelve are positive, including seven of the eight blue-collar occupations. All the negative correlations are small whereas many of the positive ones are quite large. This suggests that the seniority variable in our regressions is catching some wage differences that might have been caught by establishment size . . .[9]

In this way, establishment size may be operating as an index of labour quality through its relation with "seniority" and accumulated work experience.

It might be argued that, due to a greater division of labour, a large plant might utilize workers within a given occupation whose average skill requirements are less than those in a small plant. Other things being equal, these lower skill requirements would lead to lower wages. However, it seems unlikely that this possibility will be sufficiently important to outweigh the other effects noted, and the strong presumption is that wages are positively associated with plant size.

There is no guide in theory as to how the plant size variable, PS, should be specified. In this situation some experimentation is justified. Accordingly, PS will be either the number of non-office employees in the

---

[8] Albert Rees and George P. Shultz, Workers and Wages in an Urban Labour Market (Chicago: University of Chicago Press, 1970).

[9] Ibid., p. 189.

establishment, or the natural log of this number. The nonlinear form appears appropriate as it is possible that differences in establishment size in the range from, say, 50 to 100 employees might be expected to have a larger effect than the difference between 300 and 350 employees when factors such as qualitative differences in the work environment or approaches to hiring are considered. Besides the intuitive appeal there is some evidence that a log form of the variable will give a better fit than the simple linear form.[10]

Working Conditions

A host of job characteristics may give rise to compensating differentials. For example, such factors as the regularity and agreeability of employment can be expected to affect wages within a given occupation. Unfortunately, there are no reasonable proxy measures for these factors at the establishment or the industry level. The best that can be done is to introduce rather crude indicators of inter-industry differences in working conditions. First, there may be some justification for introducing a variable, D, that distinguishes between durable and non-durable goods industries. The former, being cyclically sensitive, may provide less stable employment opportunities and consequently a compensating differential. In addition, durable goods industries may require heavier, more arduous work, or unattractive working conditions.[11]

Similarly, the ratio of males to total employment, M, may also act as a proxy measure of working conditions with industries where the ratio is high paying higher wages. The inclusion of this variable might also be justified if it is believed that overall labour force characteristics affect industry wage rates. For example, the earnings of male workers might be lower in an industry with a large proportion of female workers if there is pressure to keep intra-establishment wage differentials as low as possible so that the terms appropriate to the main core of the work force are to a degree extended to all workers in the shop.[12]

Finally, it should be added that restricting the analysis to manufacturing industries does serve to increase the inter-establishment comparability of working conditions.

---

[10] Ibid., p. 189.

[11] Weiss, "Concentration and Labour Earnings," p. 99.

[12] Ibid., p. 99.

Employment Growth

Another possible explanation of intra-occupational wage dispersion supported by competitive market theory is based on short-run shifts in labour supply or demand. To the extent that these produce area wage differentials, they would be taken into account by the variables that pick up both regional and urban location influences. However, another market impact is possible if it is assumed that intra-industry mobility is high relative to inter-industry mobility within a given region, and that the short-run supply of labor to an industry is less than perfectly elastic. Given relative stability of labour supply functions, changes in industry employment for a given occupation will index changes in short-run labour demand and will be directly associated with changes in wage levels. Under these assumptions, one might expect establishments in industries where employment in a particular occupation has been expanding rapidly to pay higher wages than others.

There are a number of problems in specifying a variable that will adequately account for the demand shift influences described above. First, as occupational employment data is not available, estimates of the growth in total industry employment must be used as a proxy variable. Second, economic theory does not point to a precise time period for the calculation of the rate of growth of industry employment. Growth over an eight-year period has been selected on the grounds of the availability of data and because it is roughly consistent with other studies.[13] Third, simply using the rate of growth of industry employment ignores the possibility that there may be significant differences in the labour supply elasticities in various industries.[14] In short, the rate of growth of industry employment, E, is only a crude index of labour demand shifts and associated wage changes. This may not represent a fundamental problem as long as it is reasonable to expect that the unaccounted effects of labour demand shifts are uncorrelated with the variables that specify union influence.

In summary, the model will attempt to take into account the influence of competitive labour market factors by a set of variables indicating regional location, urban location, plant size, broad industry type, ratio of males to total employment, and industry employment growth. The list

[13] See, for example, ibid., p. 99.

[14] For an elaboration of this point, see: G. Rosenbluth, "Wage Rates and the Allocation of Labour," The Canadian Journal of Economics, Vol. 1, No. 3 (Aug. 1968), pp. 566-82.

TABLE 3

NON-WAGE LABOUR COSTS AS A PERCENTAGE OF GROSS PAYROLLS FOR WAGE EARNERS IN MANUFACTURING INDUSTRIES IN CANADA, 1968

Industry	Paid Absence	Premium Pay	Benefit Plans
	%	%	%
HIGHLY UNIONIZED*			
Tobacco processing and products	11.3	3.2	6.9
Rubber products	8.6	4.2	10.4
Textile products	7.1	3.6	3.5
Paper and allied industries	8.5	4.8	4.7
Printing, publishing and allied industries	8.1	4.0	2.9
Primary metals	8.8	4.1	7.5
Machinery (except electrical)	8.1	2.9	7.0
Transportation equipment	8.3	5.4	10.2
Electrical products	8.5	2.5	4.9
Unweighted average	8.6	3.9	6.4
OTHER INDUSTRIES			
Food and beverages	7.6	2.8	3.5
Leather products	5.8	1.7	2.0
Knitting mills	5.3	0.9	1.1
Clothing	6.4	0.9	1.7
Wood products	7.0	2.1	1.9
Furniture and fixtures	6.1	1.7	2.3
Metal fabricating	7.3	2.6	4.3
Non-metallic mineral products	7.3	4.2	4.5
Petroleum and coal products	11.6	5.2	7.6
Chemicals and chemical products	9.0	3.8	5.8
Miscellaneous manufacturing industries	6.8	2.7	3.8
Unweighted Average	7.3	2.6	3.5

SOURCE: D.B.S. Catalogue No. 72-510, Occasional.

*Industries in which union membership as a percentage of production worker employment exceeds 60 per cent. The union membership data used in this calculation are the estimates provided by the Canada Department of Labour and the employment estimates were derived from the Census of Manufacturing.

excludes sex and skill level as the model will be used to explain wage variations between establishments for particular sex-occupation classifications.

Omitted Variables

The model as developed above does not include many variables that competitive labour market theory suggests might be relevant for wage determination. However, the omitted variables do not pose a serious problem as long as they are uncorrelated with the variables included in the model that specify the size of the union-nonunion wage differential. If this is true, despite the misspecification, the parameter estimates will be unbiased. There is a need to determine whether there is an a priori expectation that any of the omitted variables are likely to be correlated with the measures of union influence.

There are perhaps two major problems in this respect. First, there appears to be a great deal of casual evidence that the non-wage terms of employment are positively correlated with unionism.[15] This notion is confirmed by even the crude comparisons embodied in Table 3. This relationship presents a difficulty because competitive labour market theory suggests that above average fringe benefits and working conditions should be accompanied by correspondingly lower wage rates. Accordingly, a failure to "standardize" the wage analysis for these non-wage conditions produces a downward bias in the measure of the union-nonunion wage differential. Unfortunately, nothing can be done within the context of the model to adjust for this source of bias.

The second problem stems from the possibility that employers who, for one reason or another follow a high wage policy, may be more selective in hiring and thereby obtain a higher quality work force. To the extent that labour quality is connected with plant size, this factor has been controlled for but not in any other way. The result may be that although observed intra-occupational wage differentials may be large, wage differences in terms of Marshallian efficiency units may be far less or totally absent.

The available data does not permit the introduction of labour quality variables that may explain intra-sex-occupation wage differentials. How serious is this omission?

The important empirical issue is the extent to which employers can adjust to high wages via labour quality adjustments. In fact, little is known

[15] See R. G. Rice, "Skill Earnings and the Growth of Wage Supplements," American Economic Review Proceedings (May, 1966), p. 586.

about this. Lester[16] and Reynolds[17] both have reported that wage differences are much larger than can be explained by quality differences but their evidence is non-quantitative. The most widely cited study on this point is that of Weiss.[18] He found that an initial positive relationship between earnings within an occupation and concentration disappeared when variables he described as controlling for labour quality were introduced. The implication drawn from this result by Weiss and others is that employers adjust rather completely to above market rates via quality adjustments, at least in concentrated industries.[19]

There are a number of reasons why the implications of Weiss's work for our study are not as important as first might be expected. First, a number of what Weiss has called "labour quality" variables are already taken into account in our approach or are of doubtful relevance. There are two sets of dummy regressors that take into account weeks worked in the year and hours worked per week. These "quality" factors are not relevant to our model where the dependent variable is in terms of either an hourly or weekly wage rate. Three other labour "quality" variables introduced by Weiss were regional location, urban residence, and sex. All of these factors are controlled for in our approach. Weiss controls for race but this would not appear to be a critical omission in the Ontario context. Some of the other variables introduced by Weiss have a questionable connection with the concept of labour quality as normally understood. For example, family size, recent geographic mobility, and the foreign born status of

[16] Richard A. Lester, "A Range Theory of Wage Differentials," Industrial and Labour Relations Review, Vol. 5, No. 4 (July 1952), pp. 483-500.

[17] Lloyd G. Reynolds, The Structure of Labour Markets (New York: Harper and Brothers, 1951).

[18] Weiss, "Concentration and Labor Earnings," pp. 96-117.

[19] There is a problem in interpreting Weiss's regression results. All the regression results are not reported and we are accordingly unable to follow in a detailed way what occurs when the "quality variables" are introduced. In particular we are unable to follow what is happening to the standard errors. This is important as in interpreting his results, Weiss appears to attach some weight to the "non-significance" of the coefficients relating to concentration after the labour quality variables are introduced. However, when a new independent variable is added to a regression equation and the coefficients change, it can be either because specification bias was present in the regression coefficients before the variable was added (as Weiss maintains) or because the added variable is irrelevant and the estimates after inclusion of the new variable are coming from a different distribution with larger variance but with the same mean. Weiss appears to pay little attention to this second alternative. Nevertheless, the fact that after the inclusion of the so-called labour quality variables the impact of concentration is "negative about as often as it is positive" cannot be ignored.

parents of native born members of the work force all are used. The only variables clearly connected with labour quality used by Weiss but not incorporated into our model are age and education. It is unclear whether these variables alone would have accounted for a significant proportion of the observed change in the concentration regression coefficient.

Perhaps the most interesting result for our work is that the relationship between unionism and earnings did not decline greatly when the labour "quality" variables were added to the regression equations. The implication would appear to be that measurable labour "quality" adjustments to high union wages are not very great, so that the omission of these variables is not a serious problem. Although convenient, this result in the case of unionism is disturbing. One is asked to believe that employers make labour quality adjustments rather completely when a firm follows a high wage policy as a result of concentration but not when high wages are produced by unions. Weiss offers no explanation of these anomalous results but they may be connected with the fact that a unionized employer has less flexibility in making quality adjustments after the initial hiring as, after the probationary period, dismissals can only be for "just cause."

There is additional evidence that in the case of high union wages measurable labour quality adjustments are in fact limited. As has been mentioned, Stafford[20], using the 1966 U.S. Survey of Consumer Finances, regressed earnings in particular occupations on personal characteristics, unionism, and industry. Stafford's unionism variable referred to whether the individual was or was not a union member. This specification would appear to be superior to the extent of trade union organization in the industry to which the individual belongs, the variable used by Weiss. While Weiss reported a moderate decline in the impact of unionism when personal characteristics were introduced, when this was done with Stafford's sample, union impact increased. Again the suggestion appears to be that measured quality adjustments to high union wages are not substantial.

Considering this evidence, the inability in this study to control for labour quality variation within sex-occupation categories does not appear as an overwhelming problem. The only two studies available suggest that measurable quality variations in response to high union wages are not great. For those who believe that unmeasurable quality variations may be important, the model can be interpreted as indicating the extent to which employers are forced to make partial or complete quality adjustments.

---

[20] Stafford, "Concentration and Labor Earnings: Comment," pp. 174-81.

All the control variables introduced are those that competitive labour market theory suggests as influences on wage levels. This almost universal practice can be defended on at least three grounds. First, it can be argued that in the absence of unionism, labour markets are essentially competitive, at least in the long-run. Even if imperfections do exist, they are likely to be relatively minor and not systematically related to either the presence of unions or the size of their impact on wages. Accordingly, the model can be interpreted as indicating the deviation of union wage rates from what they would otherwise be. Second, it must be recognized that it is very difficult to obtain observable measures of labour market imperfections suitable for use in wage determination equations. For example, it is difficult to introduce a variable that would in some way capture inter-establishment or inter-industry differences in the degree of employer monopsony power or deviations from maximizing behaviour. Third, even if it is believed that nonunion markets are inherently imperfectly competitive, controlling for competitive labour market influences in a model of the type used here can still be a useful exercise. In this case the model indicates deviations of union rates from an idealized normative standard. This alternative interpretation of the model should be kept in mind even though, in the discussion to follow, the empirical results are interpreted as the actual union impact on the wage structure.

Union Impact Variables

No widely accepted theory of trade union behaviour exists and no attempt at developing such a theory will be made here. Our objective is quite modest. A brief review will be made of the factors that neoclassical wage theory suggests might influence a union's wage success.[21] The list of factors discussed is of necessity short as the intent is to specify a model that can be empirically useful in explaining the inter-industry structure of wages. In such a study, much of the fine detail of considerable interest in specific situations must be omitted.

The Derived Demand Elasticity

Unions are concerned with both the employment of their members and high wages. Accordingly they can be expected to be more effective in

[21] For a discussion of the respective roles of economic versus political explanations of union behaviour see: Arthur M. Ross, Trade Union Wage Policy (Berkeley: University of California Press, 1956); Campbell R. McConnell, "Institutional Economics and Trade Union Behaviour," Industrial and Labor Relations Review, Vol. 8, No. 3 (April 1955), pp. 347-60; John T. Dunlop, Wage Determination Under Trade Unions (New York: Augustus M. Kelley, 1950).

raising wages when they face a relatively inelastic demand for labour. Marshall has made the often repeated point that the demand for one of a number of jointly demanded factors of production will be more inelastic: (1) the more essential the given item is in the production of the final product; (2) the more inelastic the supply of co-operating factors; (3) the more inelastic the demand for the final product; and (4) the smaller the fraction of total costs accounted for by the factor in question.[22] By incorporating proxy variables for at least some of these determinants into the model at least part of the inter-industry differences in union wage effectiveness may be captured.

The impact of the elasticity of product demand can be considered at two levels. If a union has been successful in organizing all the firms in an industry and the potential development of a nonunion sector is not a problem, it is the elasticity of product demand at the industry level that must be of concern. Measures of product demand elasticities for a wide range of industries are not available and accordingly this determinant cannot be considered in the model. However, in a poorly organized industry, the products of nonunion firms, actual or potential, will be close substitutes for the products of union firms and the resulting elasticity of demand for union products will be high. One would expect, therefore, a positive relationship between a union's wage gaining ability and the extent of union organization in the industry, U.

The relationship between union impact and U may not be direct or simple. It may be that some critical level of union organization in an industry must be reached before the nonunion sector ceases to be a significant threat to union power. To allow for the possibility of nonlinear effects, U will be specified as a set of binary regressors corresponding to various levels of union organization. A simple linear form will also be used.

The direct relation between labour costs as a proportion of total costs and labour demand elasticity has frequently been cited as support for the hypothesis that craft unions can be expected to be more effective than industrial unions.[23] However, it can also support the hypothesis that industrial unions in industries where production labour costs are a small proportion of total costs will be relatively more effective in raising wages than other industrial unions. The required cost information is not available

---

[22] Actually this only holds true when the elasticity of substitution is less than the elasticity of demand. See J. R. Hicks, The Theory of Wages 2nd ed. (London: Macmillan, 1963), Appendix.

[23] See, for example, Milton Friedman, Price Theory: A Provisional Text (Chicago: Aldine Publishing Co., Revised Edition 1962), p. 156.

at the establishment level. However, the ratio of production worker wages and salaries as a proportion of value added for the industry, W/VA, can be used as a proxy measure.

Market Structure

There is considerable support for the notion that the market structure that a union faces is important in determining its wage gaining ability but there is no clear concensus on the direction of the effect. On the one hand, there are those who argue that a union, given the same level of union organization in an industry, has an advantage in dealing with an oligopolist or monopolist rather than a firm in a highly competitive industry.[24] Three advantages are frequently cited. First, in non-competitive markets, well-defined patterns of price leadership involving both union and nonunion employers facilitate the maintenance and passing on of high wages in the form of high prices. Second, in non-competetive markets significant barriers to entry are likely to exist and consequently the emergence of a nonunion sector is not likely to be a serious threat. Even if new firms do enter they would normally be large and consequently subject to organization or forced to follow union rates. Finally, firms in non-competitive industries might be viewed as having a higher "ability-to-pay" in that wage increases may be absorbed by a reduction in monopoly rents.

Not all writers share the above view. Levinson[25] for example, has noted that in some highly organized and competitive industries (construction, longshoring, and trucking) unions appear to have a substantial wage impact. He argues that the prime determinant of a union's power is its ability to maintain "jurisdictional control" and that in some competitive industries unions are able to do this irrespective of how easy entry into the industry might be. It is recognized that a union's ability to maintain jurisdictional control in manufacturing industries generally goes hand in hand with a concentrated market structure. But this is viewed as being incidental. He goes on to argue that given extensive "jurisdictional

[24] See Martin Segal, "The Relation Between Union Wage Impact and Market Structure," Quarterly Journal of Economics, Vol. 78, No. 1 (Feb. 1964), pp. 100-11; Rapping, "Monopoly Rents, Wage Rates and Union Wage Effectiveness," pp. 31-47; John Kenneth Galbraith, "The Class Struggle in The Industrial State," Paper delivered to the National Tripartite Conference. Ottawa, Oct. 26-29, 1969.

[25] H. M. Levinson, "Unionism, Concentration and Wage Changes: Toward a Unified Theory," Industrial and Labour Relations Review, Vol. 20, No. 2 (Jan. 1967), pp. 198-204; ________ Determining Forces in Collective Bargaining (New York: John Wiley & Sons, Inc., 1966).

control" a union's wage gaining ability in a competitive industry may be greater than in a non-competitive industry. In these circumstances, the "resistance power" of individual competitive employers may be relatively weak (due to the strategy of divide and conquer) when compared to strongly unionized and concentrated industries where employers have the means to resist. This suggests that, if "jurisdictional control" can adequately be represented by the extent of union organization in an industry and this factor is controlled for in the analysis, there may be a negative relation between the degree of market power and union wage rates.

Rees is also skeptical about whether an imperfect product market structure increases a union's wage gaining ability.[26] He examined the extreme conditions of perfect competition and monopoly and concluded that there is no reason to believe that the elasticity of labour demand is related in a unique way to product market conditions. If anything, a monopolist can be expected to be operating on a portion of an industry's demand curve that is more elastic than that which would be applicable for a comparable competitive industry.[27]

In summary, there are a number of arguments that relate a union's wage gaining ability to product market structures. Although, popular opinion and a number of economists suggest that unions can be expected to obtain higher wages in non-competitive product markets, it is also possible that a negative relationship might exist, especially if other sources of union strength are taken into account.

The most widely used method of capturing the influence of product market structures is to introduce a measure of industry concentration, CR, as an explanatory variable. In this study, the Herfindahl concentration index based on the value of factory shipments for Canadian manufacturing industries in 1965 will be used. This single measure has been selected primarily on the grounds of its availability for a broad sample of industries and the fact that prior research has demonstrated that the particular measure of concentration selected is not critical.[28]

[26] Albert Rees, The Economics of Trade Unions (Chicago: The University of Chicago Press, 1962), pp. 82-87.

[27] Ibid., n. 2, p. 86.

[28] See Robert W. Kilpatrick, "The Choice Among Alternate Measures of Industrial Concentration," The Review of Economics and Statistics, Vol. XLIX, No. 2 (May 1967), pp. 258-60. The author concludes: "The investigation has failed to label any concentration measure as the best structural indicator of market power. The comparison of alternatives has, however, provided much evidence that the particular choice is not crucial . . . Adjustments for imports and regionalism perhaps have some use, but with present information they add little explanatory power to the unadjusted ratio."

Concentration, as a measure of market structure, is subject to some well-known limitations. First, the relationship between the number and size distribution of firms and market performance is not conceptually neat. Second, the industrial classifications used in compiling concentration ratios frequently do not coincide with the economist's conception of relevant industry definitions. Third, the use of national concentration ratios does not do justice to the actual diversity in the geographical extent of markets between industries. In particular, the existence of distinct regional or local markets and foreign trade is glossed over by national concentration ratios. Finally, the exclusive reliance on concentration does not take into account other important determinants of market performance such as product differentiation and barriers to entry.

To partially correct for these deficiencies other indicators of market position can be added. Another product market structure characteristic that may play a role is the level of barriers to entry into the industry. If the optimum firm size is large relative to the size of the industry and diseconomies of less than optimum scale operation are large, the potential entrant must either depress the market price and/or incur relatively higher costs. Accordingly, some room is provided for wages and prices to exceed competitive levels before entry appears attractive. Another consideration is that when entry into an industry must be on a relatively large scale, there is less of a change that the entrants will compete by offering relatively low wages. If large, they are likely to be resigned to the prospect of being organized or feel sufficiently threatened by the possibility of union organization that they follow the prevailing wage standards of the industry. However, if new firms can enter on a small scale, they are likely to feel these pressures less.

This discussion suggests the inclusion of a variable indicating the optimum size of firms for the industry. Unfortunately, estimates of optimum firm size for a wide range of industries are not available. As a crude approximation, the average size of establishments in an industry, AS, will be used.

Another aspect of market structure that may have an impact on a union's wage gaining ability can be directly related to establishment size, PS. If, in a particular market, there are significant economies of scale in production, it may be that unions permit the survival of smaller establishments, despite their cost disadvantage, by accepting lower wages in these establishments. A union may realize that to force a small employer to pay the same wages as the largest employer may cause the former to go out of business. Accordingly, to the extent that small firms have higher costs, the union's bargaining power may be less in these firms.

This may explain in part the common identification of "bigness" with a high ability-to-pay.

A relationship between union impact and plant size may also arise due to idiosyncrasies in industrial classification. A given "statistical" industry may be comprised of both small and large establishments. However, it may be that the large firms actually operate in their own basically non-competitive markets while the small firms operate in related but distinct markets which in many cases may be highly competitive. In such a situation, plant size may act as a proxy for market conditions.

A union's bargaining power may also vary with plant size. While large employers cannot generally maintain production during strikes (at least in manufacturing), the small employer has more opportunities for continued production. Strikebreakers may be hired or, in the extreme case, movement to a new location may be considered. Moreover, continued production is not as likely to be as critical for the small employer as his customers are bound to have temporary substitutes available. In short, even though the large employer can be expected to have greater financial resources available for withstanding a strike, this advantage may be outweighed by the small employer's more flexible position.

For any or all of these reasons, a relationship between plant size, PS, and a union's wage gaining ability may exist. Moreover, it may be that the union in a large establishment in concentrated as opposed to competitive industries has particular advantages. This suggests the use of an interaction variable involving PS and CR.

Profits

Stemming from its frequent use in collective bargaining, profits have often been cited as a variable affecting the ability of unions to win wage increases. Profits or monopoly rents have been viewed as one important component of the rather vague concept of ability-to-pay.[29] On the assumption that some, but not all, monopoly rents are absorbed in higher wages, the result is a relation between actual observed profit levels and wage rates.

The specification of the profit variable encounters at least two major difficulties. First, there is likely to be a positive correlation between profit rates and wage levels even if perfectly competitive behaviour is assumed.

[29] David G. Brown, "Expected Ability to Pay and the Inter-industry Wage Structure in Manufacturing." Industrial and Labor Relations Review, Vol. 16, No. 1 (Oct. 1962), pp. 45-62; Sumner Slichter, "Notes on the Structure of Wages," Review of Economics and Statistics, Vol. 35, No. 1 (Feb. 1950), pp. 80-91; Rapping, "Monopoly Rents, Wage Rates and Union Wage Effectiveness," pp. 341-47.

For example, given short-run inelasticities in the supplies of capital and labour, an increase (decrease) in industry product demand will lead to relatively high (low) wage and profit rates. Therefore, the profit rate for any given year may act as an index of labour demand and consequently cannot be used in isolation as a test of the hypothesis that industries with "permanent" as opposed to "transitory" monopoly rents pay high wages.

Two things can be done to meet this difficulty. Demand fluctuations can be controlled for by the inclusion of another variable. In fact, the model does contain such a variable, the rate of growth of industry employment. The second adjustment that can be made is to use average profit rates over a number of years rather than just one year. Such an average will be dominated to a greater extent by permanent as opposed to transitory forces. The data source on profits allows the use of industry average profit rates over the period 1965-68.

The second difficulty with the simple approach is that it is not at all clear that a relationship between profits and wages implies a relationship between profit rates and wage levels. Rapping illustrates this point with the following example.[30] Consider two monopolized industries with different capital/labour ratios. Industry A uses 10 units of capital and 100 units of labour, while B uses 100 units of capital and 10 units of labour. Industry A has two units of monopoly rents and Industry B has five units of monopoly rents, so that the excess rate of return in A is 20 per cent and in B, 5 per cent. If labour receives all the monopoly rents, there will be two units of monopoly rents in A for 100 units of labour and five units in monopoly rents of 10 units of labour in B. Even though the excess rate of return is higher in A than in B, the wages of this industry, ceteris paribus, would be lower. In short, the rate of return on capital invested cannot be used directly in testing the "ability-to-pay" hypothesis because it fails to take into account variations in the capital/labour ratio.

The above indicates that the most relevant conceptual profits variable is labour's share of monopoly rents per man hour. Rapping proceeds to develop a simple model which suggests that this concept can be measured by the rate of return times assets per man hour, R·A/MH, and assets per man hour, A/MH.[31] Both this specification and the more conventional rate of return, R, will be used. In both cases, the profit concept used is the rate of return on equity assets.

[30] Rapping, "Monopoly Rents, Wage Rates, and Union Wage Effectiveness," pp. 37-38.

[31] See Appendix B for an elaboration of this point.

In summary, in the model used here a union's wage gaining ability has been related to the extent of union organization, U; wages as a proportion of total costs, W/VA; industry concentration, CR; average firm size in the industry, AS; plant size (PS and PS·CR); and profits R, or R·A/MH and A/MH. All of these variables are expected to have a positive influence on a union's wage gaining ability except W/VA where the expected effect is negative. Although it is expected that CR will be positively related to union wage rates, a negative relation might be interpreted as supporting the hypothesis suggested by Levinson. All the variables, except plant size, relate to industry characteristics not available from the wage rate survey. Appendix C describes the exact specification of these variables and the supplementary data sources used.

Spillover Effect Variables

Unionism in an industry can be viewed as having a primary effect on the nonunion firms in that industry in two ways. The first effect is related to shifts in the supply and demand for nonunion labour generated by raising union wages above competitive levels. Assuming that the union and nonunion labour in an industry are substitutes in demand, the high union wage will be reflected in an increase in the demand for nonunion labour. At the same time, if the labour displaced by the high union wage is limited in the short-run to intra-industry mobility, the relative supply of labour to nonunion firms will be increased. However, both these effects (which, incidentally, operate in different directions on the wage rates of nonunion firms) are likely to be negligible in the face of a highly elastic supply of labour in the long-run.

It would appear reasonable to expect such high elasticities for the occupations being analyzed in this study as they are all found in a wide range of industries and areas. Moreover, it is important to note that short-run changes in the supply and demand schedules facing nonunion firms are only connected with a changing union-nonunion wage differential rather than one that has remained constant for some time. For these reasons, it seems likely that this type of impact of unionism on nonunion firms will only give rise to transitory short-run wage effects which can safely be ignored in our analysis.[32]

[32] This is consistent with the conclusion of Lewis, Unionism and Relative Wages in the United States, pp. 19-30.

The second effect, which appears more likely to be of some consequence, is related to the persistent threat that nonunion firms may become organized.[33] Faced with this possibility, the nonunion firm will follow a wage policy that minimizes its expected future wage costs. If we ignore all labour costs but wages and abstract from the problem of discounting future streams of wage costs to their present value, the nonunion employer can be viewed as having an expected extra wage payment (above supply price) due to the threat of unionism, defined by the expression:

$$E(\alpha) = (1-P) \cdot \alpha_n + P \cdot \alpha_u \tag{10}$$

where $E(\alpha)$ is the employer's expected extra wage payment, and P is the probability of the firm becoming organized.

In this system P is, inter alia, a function of the difference between union and nonunion wage rates, i.e.,

$$P = P(\alpha_u - \alpha_n) \quad \text{where} \quad \frac{\partial P}{\partial \alpha_u} > 0 \quad \text{and} \quad \frac{\partial P}{\partial \alpha_n} < 0.$$

From the nonunion employer's point of view, α_u is determined exogenously. The problem for the nonunion employer is to select a value, $\alpha_n{}^*$, that minimizes his expected extra wage costs, given α_u.

The first order condition for this minimum requires that

$$(1\text{-}P) - \alpha_n^* \cdot \frac{\partial P}{\partial \alpha_n} + \alpha_u \cdot \frac{\partial P}{\partial \alpha_n} = 0$$

or $$\alpha_n^* = \frac{\alpha_u}{\dfrac{-(1-P)/P}{P_e} + 1} \quad \text{where } P_e = \frac{\alpha_n^*}{P} \cdot \frac{\partial P}{\partial \alpha_n}. \tag{11}$$

This expression illustrates the reasonable propositions that a nonunion response to the threat of union organization will be greater the higher the union wage rate, the greater the probability of union organization, and the greater the effectiveness of a nonunion wage response in reducing the probability of union organization.

[33] This discussion follows closely that given by Rosen, "Trade Union Power, Threat Effects and the Extent of Organization," pp. 192-95.

In order to move from (11) to its empirical counterpart some assumptions concerning the arguments of P must be made. Apart from the size of the difference between union and nonunion wage rates ($\alpha_u - \alpha_n$), the probability of organization and perhaps the elasticity measure, P_e, are likely to vary with two critical characteristics. First, the extent of trade union organization in an industry, U, is presumably a good indicator of the willingness of employees in an industry to join unions and also the organizing ability and aggressiveness of the union or unions operating in the industry. Accordingly, nonunion employers in well-organized industries are likely to feel more threatened than their counterparts in poorly organized industries. The size of the establishment, PS, is also likely to be of some importance as there is reason to believe that union organizing costs per employee fall as plant size increases and that employees in large plants more readily join trade unions.[34] These arguments suggest that:

$$P = P(\alpha_u - \alpha_n, U, PS) \tag{12}$$

where $\dfrac{\partial P}{\partial PS}$ and $\dfrac{\partial P}{\partial U} > 0.$

Equations (11) and (12) cannot be used directly in the model as we were unable to specify a probability function that would result in a reduced form equation linear in parameters and hence subject to estimation by standard techniques. As an approximation let it be assumed that spillover effects are a varying proportion of the expected extra union wage, depending upon the extent of unionism in the industry and plant size. In particular, it is assumed that spillover effects can be approximated by:

$$\alpha_n = \alpha_{n_0} + (\alpha_{n_1} U \cdot PS + \alpha_{n_2} U)\, \alpha_u \tag{13}$$

where α_{n_1} and $\alpha_{n_2} > 0.$

At least three aspects of this specification suggest that it might be a reasonable approximation. First, spillover effects are made to vary directly with the size of the expected union impact. Where unionism has a negligible impact on wages, spillover effects will be small. Second, a high level of unionization is set as a prerequisite for large spillover effects. It is

[34] Joseph Shister, "The Logic of Union Growth," in Readings In Labour Economics and Industrial Relations, ed. by Joseph Shister, (2nd ed., New York: Lipincott, 1956), p. 53.

expected that the total spillover effects will be small for nonunion firms in poorly organized industries. Third, plant size operates as a factor determining the size of spillover effects only in conjunction with the extent of union organization. A large firm in a highly unionized industry will be subject to large spillover effects but such a plant in a poorly organized industry will not be.

The simplifying assumption of (13) permits the model to be re-written in a form that, although awkward, is suitable for estimation. Making the appropriate substitutions in (9) leads to the following:

$$W = \beta X + \alpha_{n_0} + (\alpha_{u_0} - \underline{\alpha_{n_0}})\, CA + \alpha_{u_i} Y_i \cdot CA +$$
$$(\alpha_{n_1} U{\cdot}PS + \alpha_{n_2} U)\alpha_{n_0}\ (1-CA) + (\alpha_{n_1} U{\cdot}PS + \alpha_{n_2} U)\ \alpha_{u_i} Y_i\ (1-CA) + e. \qquad (14)$$

A close examination will indicate that all the structural parameters can be identified from the regression coefficients of an equation corresponding to (14). Hence, this form of the model should permit a comparison between the wage level of nonunion establishments where spillover effects can be presumed to be absent and the average wage in union establishments. The result will be what has been defined as actual as opposed to observed union-nonunion wage differentials.

In summary, three basic models that may be used in an analysis of union-nonunion wage differentials have been outlined. In the first, Variant I, it is assumed that the impact of unionism is constant across all union establishments and also constant for all nonunion establishments. The model consists of a set of control variables (indices of location, working conditions, plant size, and employment growth) and a dummy variable indicating the presence of a union in the establishment. In the Variant II models, union impact is allowed to vary for union establishments. This is accomplished through the addition of union impact variables indicating labour demand elasticities, market structure, and profit levels. In the Variant III model, variation in the impact of unionism on nonunion wages is provided for through the addition of spillover variables. It has been argued that these can be represented by interaction terms involving the union impact variables, the extent of union organization, and plant size. In the next chapter all three models are used in an analysis of basic (unskilled) male labour rates.

CHAPTER III

THE MALE BASIC LABOUR RATE

Of all occupational wages, the male basic (unskilled) labour rate warrants an extended analysis for a number of reasons. At the pragmatic level, it has the substantial advantage of being available for all establishments in our sample and hence allows comparisons for a relatively homogeneous classification of labour across a large number of industries. In addition, the interindustry wage structure for unskilled labour is likely to be better defined than in the case of other occupations given that this type of labour is particularly exposed to labour market pressures. This occurs because the unskilled labour category in manufacturing is commonly identified as a critical port of entry for the internal labour market of the firm. Most new hires occur at this level and semi-skilled jobs are filled through promotions from the entry level jobs.[1] This characteristic also helps explain why the unskilled wage rate is often identified as a key benchmark rate in labour negotiations and wage administration programmes. As such, the basic male rate can be expected to index the general level of the wage structure for all unskilled and semi-skilled workers in the establishment. Accordingly, the result of the analysis of basic male labour rates can be interpreted as having wider application. A final advantage of the basic male rate, from the point of view of economic analysis, stems from the fact that the survey defines this rate as the lowest paid to labourers or equivalent unskilled employees. As a result the rate, to a large extent, will be unaffected by the accumulated work experience and other personal characteristics of those actually employed.

The results of the analysis of the basic male rate are reported under three headings. In the first section results for equations corresponding to the Variant I approach are reported along with a discussion of the coefficients of the control variables. In the second section, various formulations of the Variant II approach are used in order to sort out the

---

[1] This view of the labour market follows: Peter B. Doeringer and Michael J. Piore, Internal Labor Markets & Manpower Analysis (Lexington, Mass.: D.C. Heath and Company, 1971); Stuart O. Schweitzer, "Factors Determining the Interindustry Structure of Wages," Industrial and Labor Relations Review, Vol. 22, No. 2 (Jan. 1969), pp. 217-25; Robert L. Raimon, "The Indeterminateness of Wages of Semiskilled Workers," Industrial and Labour Relations Review, Vol. 6, No. 2 (Jan. 1953), pp. 180-94; Lloyd A. Reynolds, The Structure of Labor Markets (New York: Harper & Brothers, 1951), pp. 44-45.

most important determinants of union wage effectiveness. This section also contains a discussion of the implied variation in observed union-nonunion wage differentials. In the third section, results for the Variant III approach are presented. The estimated coefficients for the spillover effect variables are then used to calculate the expected nonunion wage rates free of indirect union effects. Using this result the actual mean union-nonunion wage differential is calculated.

The Variant I Model and Control Variables

The estimating procedure used throughout this study is ordinary least squares.[2] Below each coefficient, a number in parentheses indicates the absolute value of the ratio of the parameter to its standard error. A parameter will be judged to be statistically significant if the ratio exceeds 1.6. This is approximately the ninety-five per cent confidence level in a one-tailed t-test. The sample size, N, is 3422 for all the regressions reported in this chapter.

The regression result for the basic male wage rate using the Variant I model is as follows:

$$\text{(1.1)}\quad W = \underset{}{.841} + \underset{(8.7)}{.227\ R_1} + \underset{(8.0)}{.237\ R_2} - \underset{(0.3)}{.008\ R_3}$$

$$+ \underset{(4.7)}{.201\ R_4} + \underset{(11.1)}{.0800\ PS} - \underset{(1.9)}{.040\ C} - \underset{(5.1)}{.090\ D} + \underset{(24.3)}{.00881\ M}$$

$$+ \underset{(5.1)}{.000992\ E} + \underset{(10.9)}{.177\ CA.}$$

$$R_2 = .278$$
$$N = 3422$$

[2] There is some suggestion in the literature that generalized least squares should be used in estimating single equation models where establishment data is regressed against variables relating to both the establishment and to the industry of the establishment, particularly when there is more than one observation per industry in the sample. See B. Imel and P. Helmberger, "Estimation of Structure Profit Relationships with Application to the Food Processing Sector," The American Economic Review, Vol. LXI, No. 4 (Sept. 1971), pp. 614-27. However, such a procedure is beyond the scope of this study. The ordinary least squares estimates are unbiased but may not have minimum variance.

The model explains twenty-eight per cent of the variation in male basic wage rates between establishments. For a cross-section study using microeconomic data, this is not unduly low.

The results indicate that union establishments pay 18 cents per hour more than comparable nonunion establishments. In order to calculate the differential in relative terms, a nonunion wage was determined by assuming particular values for the independent variables. Assuming a non-city location in the Eastern Ontario - Lake Ontario region (the suppressed category), a non-durable manufacturing industry, and mean values for the plant size, ratio of males, and employment growth variables,[3] yields a predicted nonunion wage of $1.97 per hour. Accordingly, the implied differential in relative terms is 9.0 per cent. Making alternate assumptions about the levels of the control variables does not change this result materially.

At this point, it is convenient to discuss the estimated coefficients of the control variables and to comment on their stability in other formulations of the model. In general, the coefficients of the control variables were found to have plausible values and to be statistically significant. Moreover, they proved to be highly stable over all the regressions as additional regressors delineating variations in union impact and spillover effects are added. This suggests that the control variables, even in the simple model underlying I.1, are capturing influences on wage rates that are independent of unionism.

The coefficients of the regressors indicating region in I.1 and throughout the other regressions consistently displayed a wage advantage of approximately twenty to twenty-five cents per hour for the areas defined as 1, 2 and 4 over the rest of the province. This indicates, as expected, that the Eastern Ontario, Lake Ontario, Midwestern Ontario and Georgian Bay economic regions are low wage areas for unskilled labour.

Location in a city, C, has an unexpected significant negative influence on wage rates in I.1. This negative influence was stable throughout all the equations although often the coefficient was not statistically significant. However, the coefficients are generally quite small, being consistently in the range of -2 to -4 cents per hour.

The coefficient indicating the influence of being in the durable goods manufacturing sector also has unexpected significant negative values in I.1 and across the other regressions. These coefficients generally ranged between -1 to -9 cents per hour. The small negative effects are related to

---

[3] Unless otherwise specified, this set of assumptions is used in calculating all predicted wage rates reported in this study.

the inclusion in the model of the ratio of males, M, another indicator of working conditions. When M is omitted, the influence of D becomes positive and significant.[4]

The ratio of males, M, plays a prominent role in the regression equations. Not only does it have a large positive impact on wages, but of all the control variables, M has the largest beta-coefficient. This suggests that it accounts for a major portion of the explained variation in wage rate levels. Indeed, when M is omitted, the explanatory power of the regression equation is almost cut in half. The size of the coefficient can be gauged by comparing the expected wages for establishments in industries with selected high and low ratios of males. Using mean values for the continuous variables and reference group values for the categorical variables, but values of M equal to 50 and 100 per cent (roughly one standard deviation above and below the mean value for the sample), the expected wage rates for nonunion establishments are calculated to be $1.75 and $2.19 respectively. This represents a relative difference of twenty-five per cent. In addition, this large effect is quite stable. The regression coefficients of M decrease only slightly in other regressions as variables delineating union impact are added. The coefficient varies in the range of .0088 to .0069, and is consistently statistically significant.

This prominent role for the ratio of males variable is somewhat surprising given its rationale as a crude index of working conditions. At a later point in this chapter, an alternate interpretation of this variable is considered.

Employment growth, E, has its expected positive influence on wages. Although statistically significant, the coefficients are generally small. Using the previously described assumptions for the values of the other control variables and industry employment growth of zero and eighty per cent[5] leads to a difference in expected wages of only four per cent. Although declining somewhat in size, the coefficients consistently remained statistically significant as additional variables were added to the equations.

---

[4] When M is omitted, the Variant I model regression result is as follows:

$$W = 1.497 + \underset{(8.0)}{.227\,R_1} + \underset{(7.7)}{.248\,R_2} - \underset{(0.9)}{.028\,R_3} + \underset{(6.4)}{.195\,R_4}$$

$$+ \underset{(8.7)}{.0684\,PS} - \underset{(3.5)}{.077\,C} + \underset{(7.4)}{.124\,D} + \underset{(2.9)}{.000608E} + \underset{(11.6)}{.204\,CA}$$

$$R^2 = .153.$$

[5] This range is approximately one standard deviation above and below the mean value of forty-two per cent.

Plant size, PS, the final control variable to be considered here, has its expected positive influence on wages. Some indication of its importance can be gauged from the following tabulation, based on equation I.1, which shows plant size wage differentials relative to a plant with twenty non-office employees.

Non-Office Employees	Wage Differentials (Cents per hour).
50	7
100	13
200	18
500	26
1000	31

In equation I.1, the plant size variable is specified as the natural log of the number of non-office employees. In regressions where the specification was simply the number of non-office employees, the calculated differentials proved to be considerably smaller.[6] However, in terms of the goodness of fit, the log specification in various equations is clearly superior to the definition involving the absolute number of non-office employees. For this reason, the log specification was selected for use.

Unlike the other control variables, plant size also appears as a determinant of union impact and spillover effects in the Variant II and III models. Consequently, its regression coefficient often changes appreciably in the various regressions. For this reason, the regression coefficients for this control variable are consistently reported in the remainder of this chapter while others are not.

The Variant II Model -- The Role of Union Impact Variables

In the Variant II approach, union-nonunion wage differentials are allowed to vary with industry and establishment characteristics. This involves the addition to the regression equations of a number of variables, all of which are interacted with the collective agreement dummy variable. Some of the results using this approach are presented in Table 4. In this set of regressions the size of the union-nonunion wage differential is related to

[6] For the Variant I model, the plant size coefficient was .000144 with a t-ratio of 8.5 and R^2 = .267. Using this result gives a wage differential of seven cents per hour for a plant with 500 employees relative to a plant with twenty employees. The comparable differential in the log specification is twenty-six cents per hour.

the extent of industry unionism, U; wages as a proportion of value-added, W/VA; industry concentration, CR; the average firm size in the industry, AS; and plant size, PS. The sequence of the addition of variables in the table corresponds roughly to our a priori confidence with respect to their role in influencing a union's wage gaining ability.

Industry Unionism

The extent of industry unionism, measured in per cent, displays a strong influence on wage rate levels. Even in the presence of other union impact variables, the coefficient remains large and statistically significant. Some notion of the size of the effect can be gained by considering the

TABLE 4

VARIANT II: REGRESSION RESULTS FOR MALE BASIC LABOUR – EQUATIONS II.1 TO II.7

VARIABLES	II.1	II.2	II.3	II.4	II.5	II.6	II.7
PS	.0605 (8.2)	.0591 (8.0)	.0608 (8.4)	.0453 (3.8)	.0453 (3.8)	.0453 (3.8)	.0454 (3.8)
CA	-.137 (3.8)	-.107 (2.9)	.361 (6.5)	.275 (3.6)	.295 (3.4)	.294 (3.4)	-.067 (0.4)
U·CA(10^{-2})	.533 (9.7)	.415 (6.7)	.403 (6.6)	.388 (6.3)	.388 (6.3)	.386 (6.1)	.968 (4.4)
CR·CA(10^{-2})		.528 (4.0)	.377 (2.9)	.373 (2.9)	.184 (0.4)	.198 (0.4)	.973 (1.8)
CR·U·CA(10^{-2})							-.0136 (2.3)
AS·CA(10^{-2})						.000390 (0.1)	.000400 (0.1)
W/VA·CA(10^{-2})			-1.228 (11.2)	-1.234 (11.3)	-1.234 (11.4)	-1.234 (11.3)	-.396 (1.1)
W/VA·U·CA(10^{-2})							-.0128 (2.5)
PS·CA				.0245 (1.6)	.0202 (1.1)	.0203 (1.1)	.0155 (0.9)
PS·CR·CA(10^{-2})					.0407 (0.5)	.0370 (0.4)	.0920 (0.9)
R^2	.297	.301	.326	.326	.326	.326	.328

change in union wage rates associated with a change in the level of industry unionism from 35 to 90 per cent. This shift, which covers the predominant range for this variable found in the sample,[7] results in an increase in wages of 22 cents per hour.[8] This is a substantial increase being greater, in fact, than the effect of the union status of the establishment observed in the Variant I model. Unions that are successful in organizing extensively in an industry do appear to derive substantial additional bargaining power.

It is not clear that the linear form of the variable used in the regressions reported in Table 4 is the most appropriate. As mentioned in Chapter II, it may be that some critical level of union organization in an industry must be reached before a union obtains a substantial wage advantage. To see if this is the case, regressions were also tried in which the industry unionism variable was represented by a set of dummy regressors that distinguish five different levels of industry-unionism. In these equations, no marked threshold effects were found to be present. With one minor exception, it was found that the union-nonunion wage differential increases approximately linearly with movement to higher levels of industry unionism.[9] Further, a comparison revealed that there was no appreciable difference in the fit associated with the two specifications, and that the implied variation in the union-nonunion wage differential over the range U = 35 to U = 90 was the same in the two cases. For these reasons, the simpler linear form of the variable was considered acceptable.

---

[7] U exceeds 90 in 6 per cent of the establishments and is less than 35 in 12 per cent of the establishments.

[8] This result was derived using the coefficient of equation II.3. Equations II.4, II.5, and II.6 give virtually the same answer.

[9] In a Variant II regression equation that also included AS, CR and W/VA as union impact variables, the coefficients of the dummy industry unionism regressors were as follows:

$$\underset{(1.8)}{-.128\, U_{25\text{-}49}\cdot CA} + \underset{(0.2)}{.015\, U_{50\text{-}74}\cdot CA} + \underset{(0.9)}{.067\, U_{75\text{-}89}\cdot CA}$$

$$+ \underset{(2.0)}{.129\, U_{90\text{-}100}\cdot CA}$$

The subscripts indicate the range of industry unionism in per cent covered by the regressors. The category of industry unionism less than 25 per cent enters the intercept term. Note that the only departure from an approximate linear trend is the suppressed category. The implied union-nonunion wage differential for industries with less than 25 per cent industry unionism is almost as large as the differential for industries where the extent of industry unionism is from 50 to 74 per cent. This anomaly may simply reflect statistical irregularities as only two per cent of the union establishments in the sample fall in the suppressed category. The explained variance in the above equation is .326 or the same as II.4 which uses a simple linear form for the industry unionism variable.

Product Market Concentration

Product market concentration is also associated with high union wages, even when other sources of union strength are taken into account. The coefficients of CR·CA are positive and significant except when other interaction variables involving CR appear in the equation. But the effect is small. A change in CR·CA from 1 to 25, the predominant range for this variable found in the sample,[10] leads to only a 9 cents per hour increase in union wage rates. These regression results are noteworthy for, despite the well-known positive correlation between industry unionism and product market concentration, the low level of aggregation used in the analysis has made it possible to separate quite clearly the effects of each of these variables. The implication is that industry unionism plays a more prominent role than product market concentration in influencing a union's wage gaining ability.

In equation II.7, there is evidence of interaction effects between these two variables. Both the coefficients of CR·CA and CR·U·CA are statistically significant. As the coefficient of the latter regressor is negative, product market concentration appears as a more important factor in raising union wage rates in weakly rather than strongly organized industries. A close examination of the coefficients reveals that the negative coefficient of CR·U·CA is relatively small in absolute value. Throughout the predominant range of values for the industry unionism variable, the overall effects of product market concentration remain positive. At very high levels of U·CA, the effects of CR are negligible. This implies that if a union is successful in completely organizing an industry, the product market structure plays no role in influencing its wage gaining ability. In any case, as the effects of product market concentration are non-negative throughout the range of feasible values for industry unionism, these results contradict Levinson's hypothesis which suggests that negative effects will appear when differences in levels of industry unionism are taken into account.[11]

Wages as a Proportion of Value-Added

The empirical results confirm the importance, from labour's point of view, of being unimportant. Throughout the regressions reported in Table 4, total wages as a proportion of value-added exerts a strong negative influence on union wage rates, which, of course, implies that where labour

---

[10] Values of CR·CA exceed 25 in 7 per cent of the establishments in the sample and are below 1 in 8 per cent of the establishments.

[11] See *supra*, p. 64.

costs are low, union wage rates are high. The size of the effects are quite large. Lowering labour costs from 50 to 20 per cent of value-added,[12] increases union wage rates by 37 cents per hour. This is the largest wage change associated with variation in any of the union impact variables over its predominant range.

Equation II.7 suggests the presence of interaction between W/VA and U, even though large standard errors make the separation of effects difficult. It does seem that highly unionized industries are in a particularly advantageous position for exploiting low labour costs.

Average Plant Size

In all the regressions containing AS its coefficients were found to be very small and far from being statistically significant. Using selected high and low values for the variable leads to virtually no change in the predicted wage rate.[13] The relevance of this variable can clearly be questioned. Either it is a poor proxy for barriers to entry or this factor plays a minor role once product market concentration and industry unionism are taken into account.

Plant Size

The effects of introducing plant size as a union impact variable can be seen in equations II.4 through II.7. The coefficient of the plant size control variable, already reduced by the inclusion of the other union impact variables, falls even further in these equations. But in all cases, it remains positive and retains its statistical significance. Both plant size union impact variables have their expected positive influence. When PS·CA is added by itself, its coefficient is just barely statistically significant. When the interaction variable, PS·CR·CA, is added as well, its coefficient and the coefficients of CR·CA and PS·CA all fail to show statistical significance, although they do remain positive. A strong interaction effect is not apparent. However, the evidence can be read as giving weak support for a small positive effect of plant size on a union's wage gaining ability.

In other words, while both union and nonunion plants have wages that are positively related to plant size, the effects are somewhat larger in union plants. The coefficients of II.5 indicate that the wage differential for a 500

[12] These values for W/VA·CA are representative of the predominant range found in the sample as W/V·CA exceeds 50 for 10 per cent of the establishments in the sample and are below 20 in 7 per cent of the establishments.

[13] For example, using values of 20 and 250, and either the coefficient in II.6 or II.7, leads to no change in the predicted wage rate. AS exceeds 250 for 9 per cent of the establishments and is below 20 for 10 per cent of the establishments.

employee plant relative to a twenty employee plant is 7 to 9 cents per hour greater for union establishments than for nonunion establishments, depending upon the level of product market concentration.

Profits

Table 5 presents regression results based on the two alternative specifications of the profit variable discussed in Chapter II. First, consider the equations involving R and R·U (II·8 through II·12). When the profit regressors represent the only union impact variable included (II.8 and II.9), their effect is large and statistically significant. However, as other union impact variables are added the role of profits declines markedly, and the regression coefficients lose their statistical significance. Moreover, a comparison of II.3 and II.11 indicates that when other union impact variables are included, the increase in the explanatory power of the model due to the addition of the profit regressors is very small, although it does

TABLE 5

VARIANT II: REGRESSION RESULTS FOR MALE BASIC LABOUR – EQUATIONS II.8 TO II.15

VARIABLES	II.8	II.9	II.10	II.11	II.12	II.13	II.14	II.15
PS	.0805 (11.2)	.0633 (8.7)	.0614 (8.3)	.0613 (8.4)	.0454 (3.8)	.0714 (9.7)	.0645 (8.7)	.0602 (8.2)
CA	.00125 (0.0)	.0414 (1.2)	–.130 (1.3)	.319 (3.1)	.248 (2.0)	.117 (6.3)	.152 (7.7)	.494 (6.9)
U·CA(10^{-2})			.258 (1.9)	.251 (1.8)	.232 (1.7)			.214 (2.6)
CR·CA(10^{-2})				.361 (2.8)	.226 (0.5)			.419 (3.1)
W/VA·CA(10^{-2})				–1.141 (10.3)	–1.147 (10.3)			–1.252 (10.8)
PS·CA					.0220 (1.2)			
PS·CR·CA(10^{-2})					.0280 (0.3)			
R·CA(10^{-2})	1.760 (5.5)	–1.646 (3.5)	.137 (0.1)	.0270 (0.0)	.0036 (0.0)			
R·U·CA(10^{-2})		.0509 (9.9)	.0286 (2.2)	.0166 (1.3)	.0169 (1.3)			
R·A/MH·CA(10^{-2}						.0184 (2.2)	.141 (3.1)	.0413 (0.9)
A/MH·CA(10^{-2})						.135 (1.5)	.695 (1.4)	–.320 (0.6)
R·A/MH·U·CA(10^{-2})							.00212 (3.8)	.000687 (1.2)
A/MH·U·CA(10^{-2})							.00816 (1.4)	.00295 (0.5)
R^2	.284	.304	.305	.329	.329	.288	.297	.328

pass the F-test at the ninety-nine per cent confidence level.

Despite the lack of statistical significance of the coefficients of the profit regressors, the implied effects are not negligible. This is indicated in the following tabulation which illustrates the variation in union wage rates in cents per hour associated with the profit variable regressors in II.11 assuming selected high and low values for R[14] and U. The tabulation also illustrates the critical role of interaction effects. At low levels of industry unionism, profits have only a small effect on union wage rates, while at high levels of industry unionism, the effect is more substantial. The implication appears to be that only strong unions are in a position to take advantage of a profitable product market environment.

R \ U	25	60	90
6	2	6	9
14	6	14	21

In equations II.13 to II.15 the profit variable specification involving A/MH is used. Although this specification is theoretically superior to the simpler formulation just described, the regression results are poorer. First of all, the regressions using this specification have a poorer fit. Second, although in both cases the individual regression coefficients are not statistically significant when other union impact variables are also included, the size of the effects based on the A/MH specification are quite small and, in certain circumstances, opposite in direction from what would be expected. This is illustrated in the following tabulation of union wage rate variation associated with the four profit variable regressors of II.15, assuming high and low levels for R and U, and the mean value of A/MH for the sample (20).

R \ U	25	60	90
6	-8	-3	1
14	-12	-3	5

[14] Rates of return in the range 6 to 14 per cent are representative of those found in the sample. Only 7 per cent of the establishments in the sample have a rate of return less than 6 per cent and only 8 per cent of the establishments have a rate of return greater than 14 per cent.

At the mean level of industry unionism, variations in profit rates have virtually no effect on union wage rates, while at high levels of industry unionism the effect is positive. However, at low levels of industry unionism increases in observed profits have a negative effect, contrary to what might be expected. This same pattern is observed when higher values for A/MH are assumed, but both the positive and negative effects are more pronounced.

The perverse negative effects do not arise because the coefficients of the profit regressors have the "wrong" signs. As explained in Appendix B, the only unambiguous prediction concerning the sign of the individual regression coefficients derived from theory involves the regressor R·A/MH U·CA. The theory predicts a positive coefficient and this is found in II.15. The theory also suggests that if the coefficients of R·A/MH·U·CA and A/MH·U·CA have the same sign, then the coefficients of R·A/MH·CA and A/MH·CA must also have the same sign. This condition is satisfied in II.15.

In general, the specification of the profit variable involving A/MH performs poorly in terms of both fit and the direction of the effects. These problems, perhaps related to substantial measurement errors, suggest the abandonment of this specification in favour of the more straightforward approach.

The results for the Variant II model may be summarized as follows. Union wage rates do vary systematically with a number of variables expected to influence a union's wage gaining ability. Foremost among the variables are the extent of industry unionism and wages as a proportion of value-added. In all formulations, the coefficients of these variables are statistically significant and large. Concentration also consistently displayed a significant influence but the size of its effects is somewhat smaller.

Other union impact variables are not as prominent. Although the direction of these effects is generally as expected, the coefficients of these variables (PS, PS·CR,R, and R·U) are often statistically insignificant and quite small. Their addition to equations including the previously mentioned variables affects R^2 by only a small amount. However, the only variable whose relevance can clearly be questioned is AS.

In point of fact, alternate formulations involving variables in addition to CR, U, and W/VA, appear to affect the results very little. This is illustrated in the following tabulation of expected nonunion and union

wage rates derived from four of the regression equations. The union impact variables included in these equations are:

II.3 – CR, W/VA, U
II.5 – CR, W/VA, U, PS, PS·CR
II.7 – CR, W/VA, U, PS, PS·CR, AS, CR·U, W/VA·U
II.12 – CR, W/VA, U, PS, PS·CR, R, R·U.

For each equation, the nonunion wage plus three expected union wage rates are presented. The union wages are calculated assuming, alternately, mean, high and low values for the union impact variables. The extreme values are the same as those used previously in analyzing the results. However, the mean value for PS was assumed in all the calculations in order to maintain comparability with the nonunion rate. As for control variables, mean values were used for the continuous variables and reference group values for the categorical variables.

TABLE 6

EXPECTED UNION AND NONUNION WAGE RATES FOR MALE BASIC LABOUR BASED ON EQUATIONS II.3, II.5, II.7, AND II.12

	II.3	II.5	II.7	II.12
	$ per hour	$ per hour	$ per hour	$ per hour
Nonunion:	1.92	1.92	1.92	1.92
Union Assuming Impact Variables With:				
Mean Values*	2.10	2.10	2.11	2.10
High Values	2.50	2.49	2.51	2.54
Low Values	1.82	1.82	1.84	1.80

*The values used in the calculations are as follows:

	Mean	High	Low
U	57.6	90	35
CR	7.8	25	1
W/VA	36.1	20	50
AS	115	250	20
R	10.0	14	6

The similarity of results for each formulation is apparent from Table 6. Even when extreme values for the union impact variables are assumed, all the formulations give virtually identical predicted results. For this reason, there would appear to be no substantial loss involved with the use of the simplest equation with only three union impact variables.

The Union-Nonunion Wage Differential in Terms of Employees

In the Variant II equations, the inter-establishment union-nonunion wage differential, evaluated at mean values for the union impact variables, remains just under ten per cent. However, the studies reviewed in Chapter I analyzed the differential between union and nonunion workers rather than establishments.[15] These two measures will differ if the number of workers in a given occupation in an establishment is positively correlated with the determinants that increase a union's wage gaining ability.

In order to arrive at an estimate of the differential in terms of employees, the expected wage rate for union establishments in each industry was calculated assuming the same set of mean or arbitrary values for the control variables and industry values for the union impact variables. The next step is to weight these expected wage rates by the number of employees affected. Unfortunately, the number of employees paid male basic rates is not available. As an approximation, an estimate of the total number of non-office male union employees[16] was used as the weighting factor in the calculations. This procedure yielded an overall employee weighted average union wage rate of $2.19. This increases the union-nonunion differential in relative terms to 14 per cent.

This calculating procedure can also be used to arrive at the distribution of male union employees by the size of the observed union-nonunion wage differential. This distribution, appearing in Table 7, shows a high concentration about the mean value of twenty-seven cents. Some 50 per cent of the union employees in the unskilled male labour category receive wages that exceed those of comparable nonunion employees by some twenty to twenty-nine cents per hour. On the other hand, it would appear that very few unskilled union members receive as much as sixty cents per hour in excess of comparable nonunion workers or less than ten cents above nonunion rates.

[15] The only exception is the study by Bailey, King and Schwenk, "Wage Differentials," where an inter-establishment differential is also reported.

[16] Calculations were made using the number of union employees times the ratio of males in total employment in each industry.

TABLE 7

PERCENTAGE DISTRIBUTION OF MALE UNION EMPLOYEES BY THE SIZE OF OBSERVED UNION-NONUNION WAGE DIFFERENTIALS FOR MALE BASIC LABOUR BASED ON EQUATION II.3

DIFFERENTIAL (cents per hour)	UNION EMPLOYEES (per cent)
Less Than –9	0.7
–9 to –1	2.9
0 – 9	7.0
10 – 19	9.9
20 – 29	51.0
30 – 39	12.1
40 – 49	4.6
50 – 59	10.6
60 – 69	0.5
70 – 79	0.7

The Effect of Omitting the Ratio of Males Variable

In the discussion of control variables, it was pointed out that the ratio of males has a surprisingly large effect on wages given its rationale as a crude index of working conditions. This large effect raises the possibility that a low ratio of males might be more appropriately considered a result rather than a cause of low wages. Consider two employers with identical technologies which utilize both male and female labour. For one reason or another, one employer has a low wage structure. Further, assume that for both employers intra-establishment differentials are kept below those prevailing in the market due to traditional views of equity that restrict the size of wage differences for those working side by side. If the low wage employer just pays the market wage for females, the wages he is willing to pay to males will be low and he will be unable to attract all the males he would wish to hire. The ratio of females in employment will be high. On the other hand, if the high wage employer pays the market wage or above for males, the low sex differential in the establishment will result in the employer wishing to utilize a high ratio of males, and his high wage structure enables him to do so. In this way, the level of the wage structure of the firm, however established, is indexed by the ratio of males. If this type of relationship is important, the inclusion of the ratio of males

variable in the wage determination equations is incorrect and, most likely, biases downwards some of the coefficients of the other variables included in the equations.

When the ratio of males, M, is omitted as a control variable from the Variant I form of the model little change is apparent. The union-nonunion wage differential increases modestly from 17.7 to 20.4 cents per hour (see Table 8). Yet in the Variant II regressions the changes are noticeable. As suspected, the effect of omitting M is to increase the absolute size of all the coefficients of the union impact variables. Notice the large plant size

TABLE 8

VARIANT I AND II: REGRESSION RESULTS FOR MALE BASIC LABOUR – OMITTING THE RATIO OF MALES AS A CONTROL VARIABLE

VARIABLES	I.2	II.16	II.17	II.18
PS	.0683 (8.7)	.0456 (5.9)	.0181 (1.4)	.0180 (1.4)
CA	.204 (11.6)	.448 (7.7)	.286 (3.1)	.052 (0.3)
$U{\cdot}CA(10^{-2})$		.517 (8.0)	.489 (7.5)	.864 (3.7)
$CR{\cdot}CA(10^{-2})$		.528 (3.9)	.607 (1.3)	1.443 (2.5)
$CR{\cdot}U{\cdot}CA(10^{-2})$				–.0132 (2.1)
$AS{\cdot}CA(10^{-2})$				.00341 (0.9)
$W/VA{\cdot}CA(10^{-2})$		– 1.621 (14.2)	–1.628 (14.3)	–1.112 (3.0)
$W/VA{\cdot}U{\cdot}CA(10^{-2})$				–.00771 (1.4)
PS·CA			.0456 (2.4)	.0412 (2.2)
$PS{\cdot}CR{\cdot}CA(10^{-2})$			–.0193 (0.2)	.0117 (0.1)
R^2	.153	.241	.243	.244

effect for union establishments appearing in equations II.17 and II.18. The implication is that union scales may be subject to considerably more variation than previously indicated. This is confirmed when the distribution of union effects by employees based on equation II.17 is calculated. Moreover, this distribution has a mean union-nonunion wage differential in terms of employees of 40 cents per hour. In relative terms, the differential is 22 per cent. This is considerably larger than the previously reported differential calculated from regressions that include M as a control variable.

The Variant III Model – Spillover Effects

As discussed in Chapter II, this approach involves a three equation system in which union effects on nonunion wages as well as union wages are allowed to vary. Equation (3), it will be recalled, is:

$$W = \beta X + \alpha_u CA + \alpha_n(1-CA) + e.$$

Based on the results reported so far it appears reasonable to assume that variations in union impact can be adequately captured by:

$$\alpha_u = \alpha_{u_0} + \alpha_{u_1} CR + \alpha_{u_2} U + \alpha_{u_3} W/VA \qquad \alpha_{u_1}, \alpha_{u_2} > 0, \alpha_{u_3} < 0. \tag{15}$$

For the moment also assume a shortened version of (13), namely:

$$\alpha_n = \alpha_{n_0} + \alpha_{n_1} U{\cdot}PS\ \alpha_u \qquad \alpha_{n_1} > 0. \tag{16}$$

These three equations reduce to:

$$W = \beta X + \alpha_{n_0} + (\alpha_{u_0} - \alpha_{n_0}) CA + \alpha_{u_1} CR{\cdot}CA + \alpha_{u_2} U{\cdot}CA + \alpha_{u_3} W/VA{\cdot}CA$$
$$+ \alpha_{n_1} \alpha_{u_0} U{\cdot}PS(1-CA) + \alpha_{n_1} \alpha_{u_1} CR{\cdot}U{\cdot}PS\,(1-CA) + \alpha_{n_1} \alpha_{u_2} U^2{\cdot}PS(1-CA)$$
$$+ \alpha_{n_1} \alpha_{u_3} W/VA{\cdot}U{\cdot}PS\,(1-CA) + e.$$

In a regression equation corresponding to this formulation, the last four regressors capture variation in union spillover effects. From the regression coefficients all the structural parameters can be identified. In fact, α_n, is overidentified.

The regression results for an equation corresponding to (17) are reported in Table 9 as III.1. This table also contains results for the comparable Variant II regression equation and an alternative formulation of the spillover effect model based on $\alpha_n = \alpha_{n_0} + (\alpha_{n_1} U{\cdot}PS + \alpha_{n_2} U)\,\alpha_u$. First, considering III.1 in relation to II.11, it is evident that the addition of the spillover effect regressors adds significantly to the explanatory power

of the model. Even though these regressors assume nonzero values for less than one-half of the establishments in the sample, the reduct on in the residual sum of squares between II.3 and III.1 does pass the F-test at the 95 per cent confidence level. Moreover, three out of the four spillover effect regressors have statistically significant coefficients with the proper sign. Using these coefficients and the related union impact variable coefficients yields three possible estimates of α_{n_1}.[17] These estimates are .00558, .00306 and .00547; all of roughly the same magnitude as the underlying model suggests.

However, an examination of the coefficient of the regressor U·PS (1-CA) reveals a trouble spot. The coefficient is negative, and, although it does not pass our test of statistical significance, the t-ratio in absolute terms is rather large. This would suggest that α_{u_0} is negative which, in conjunction with the positive value for the coefficient of CA, in turn suggests that α_{n_0} is negative and large in absolute value. As the overall contribution of the spillover effect regressors, evaluated at sample mean values of the variables, is negative, the results suggest negative total spillover effects. This interpretation, which implies that unionism causes nonunion employers to pay below the supply price of labour, does not make economic sense and is rejected.

The reason for this result is not clear but it probably arises because the functional forms used are only approximations for the true (but unknown) forms. In this case the intercept values, and in particular α_{n_0}, may contain a large component attributable to the error of approximation. If this is true, no operational interpretation can be placed on the estimates of these intercept parameters.[18] In particular, the results do not necessarily imply that spillover effects are negative.

In this regard notice should be taken of what occurs to the coefficients of the union effect variables when the spillover effect variables are added. In moving from II.3 to III.1 all the coefficients of the union impact variables remain basically the same but the coefficient of CA drops by .10. This drop in level is associated with the overall negative effect of the four spillover effect regressors evaluated at mean values.

Consider now the results for III.2 based on the assumption that $\alpha_n = \alpha_{n_0} + (\alpha_{n_1} U{\cdot}PS + \alpha_{n_2} U)\,\alpha_u$, that is, that spillover effects depend on the level of industry unionism as well as industry unionism and plant size

---

[17] It should be noted that the "indirect" least squares estimates of the parameters in the Variant III equations will, in general, be biased but consistent due to the non-linear restrictions on the parameters.

[18] See Roa and Miller, Applied Econometrics, pp. 3-6.

TABLE 9

VARIANT III: REGRESSION RESULTS FOR MALE BASIC LABOUR – EQUATIONS III.1 TO III.4

VARIABLES	II.2	III.1	III.2	III.3	III.4
PS	.0608	.0708	.0689	.0581	.0559
	(8.4)	(8.3)	(7.7)	(2.2)	(2.0)
CA	.361	.260	.239	.203	.212
	(6.5)	(4.0)	(3.1)	(1.9)	(1.8)
U·CA(10^{-2})	.403	.384	.387	.386	.386
	(6.6)	(6.3)	(6.4)	(6.3)	(6.3)
CR·CA(10^{-2})	.377	.378	.378	.378	.204
	(2.9)	(3.0)	(3.0)	(3.0)	(0.5)
W/VA·CA(10^{-2})	–1.228	–1.292	–1.292	–1.291	–1.291
	(11.2)	(11.9)	(11.9)	(11.9)	(11.9)
PS·CA				.0134	.00972
				(0.4)	(0.3)
PS·CR·CA(10^{-2})					.0374
					(0.4)
U·PS(1-CA) (10^{-3})		–.669	–.790	–1.133	–1.232
		(1.3)	(0.7)	(1.6)	(1.5)
CR·U·PS(1-CA) (10^{-3})		.0211	.0186	.0202	.0271
		(2.8)	(0.6)	(2.7)	(1.0)
W/VA·U·PS(1-CA) (10^{-3})		–.395	–.0361	–0395	–.395
		(7.0)	(1.6)	(7.0)	(7.0)
U^2·PS(1-CA) (10^{-3})		.0210	.0231	.0219	.0218
		(4.2)	(1.6)	(3.5)	(3.5)
U·PS^2(1-CA) (10^{-3})				.0881	.115
				(0.9)	(0.8)
U·PS^2·CR(1-CA) (10^{-3})					–.00167
					(0.3)
U(1-CA) (10^{-3})			–.211		
			(0.0)		
U^2(1-CA) (10^{-3})			–.00204		
			(0.0)		
U·CR(1-CA) (10^{-3})			.0078		
			(0.1)		
W/VA·U(1-CA) (10^{-3})			–.0140		
			(0.2)		
R^2	.326	.344	.345	.345	.345

interacted. In this equation the added spillover effect regressors improve the explanatory power of the model very little. In fact the null hypotheses of $\alpha_{n_2} = 0$ cannot be rejected at the 95 per cent confidence level.[19] In addition, the coefficients of the previously used spillover effect regressors involving U·PS are not greatly altered in the new equation while the coefficients of the added spillover effect regressors are, as can be seen from close inspection, all very small. But, the presence of the added regressors does affect the standard errors. Only two of the original spillover effect coefficients retain their statistical significance and none of the new coefficients is close to being significant. All these results, considered together, suggest that the simpler formulation is adequate. It would seem that both a high level of industry unionism and a relatively large plant size are necessary for spillover effects.

Alternative formulations of the union impact variables were also tried. In III.3, reported in Table 9, PS·CA is added to the list of union impact variables and in III.4 the interaction variable, PS·CR·CA, is also added. Neither equation reveals a substantial change in the results. Although these added union impact variables have their expected positive influence, the coefficients are small and not statistically significant. In III.3 the added spillover effect regressor $U{\cdot}PS^2(1\text{-}CA)$ has the expected positive sign but is not statistically significant. In III.4, one of the two added spillover effect regressors has the wrong sign and neither is statistically significant. Comparing III.3 and III.4, with III.1, the reduction in the residual sum of squares associated with the increase in the number of regressors is not statistically significant at the ninety-five per cent level for either equation. It would seem that the relevance of the added regressors is open to question.[20]

At this point, the possible size of spillover effects is considered. This cannot be done directly through an examination of the estimates of the structural parameters because of the previously mentioned specification problem. While the regression coefficients indicate the required positive values for α_{n_1}, the implied value for α_{n_0} is not acceptable. This probably arises because the assumptions underlying the spillover effect equation, (16), only crudely approximate the true functional form. Nonetheless, it is still instructive to examine the pattern of predicted nonunion wages that

[19] This, of course, involves testing the null hypothesis that the coefficients of the four added spillover effect regressors are all equal to zero. The F-value turns out to be 0.1, far below the critical value of 2.37.

[20] Identical results were obtained in regression equations corresponding to the extended form of the model with

$$\alpha_n = \alpha_{n_0} + (\alpha_{n_1} U{\cdot}PS + \alpha_{u_2} U)\,\alpha_u.$$

flows from the regression coefficients to determine whether the results are generally consistent with the existence of spillover effects.

In Table 10, expected union and nonunion wages are compared assuming high, mean and low values for the union impact variables. Expected wages calculated from three of the regression equations are presented. The pattern that emerges can be interpreted as being consistent with the existence of spillover effects. In circumstances that favour a high union impact and also large spillover effects on nonunion employers, including a high level of industry unionism and a large plant size, both union and nonunion wages are high. In fact, the two expected wages are virtually identical. But in circumstances where union impact and spillover effects are moderate, union and nonunion wages diverge substantially. At the other extreme, where union impact is low and it can be supposed that union employers do not pay wage premiums substantially in excess of the supply price of labour, both union and nonunion wages are at roughly the same level. In fact the expected nonunion wages exceed the expected wages for union employees.[21]

TABLE 10

EXPECTED UNION AND NONUNION WAGE RATES FOR MALE BASIC LABOUR BASED ON EQUATIONS III.1 AND III.4

	III.1		III.4	
Union Impact	**Union**	**Nonunion**	**Union**	**Nonunion**
High	2.53	2.52	2.54	2.51
Mean	2.08	1.86	2.08	1.86
Low	1.70	1.75	1.71	1.76

NOTE: The values used in the calculations are as follows:

	High	Mean	Low
CR	25	7.76	1
U	90	57.56	35
PS	5	4.13	3
W/VA	20	36.07	50

[21] This last result is of no particular significance and may simply reflect that extreme values have been used in the calculations.

The consistency of the regression results with the existence of spillover effects can also be considered by examining the expected changes in nonunion wages as the extent of industry unionism increases. Taking the partial derivative of (16) with respect to U gives:

$$\frac{\partial \alpha_n}{\partial U} = \alpha_{n_1} PS\ \alpha_u + \alpha_{n_1} U \cdot PS \cdot \frac{\partial \alpha_n}{\partial U}. \qquad (18)$$

This expression suggests that, if spillover effects are operating, an increase in U will lead to an increase in nonunion wages. Furthermore, the increase will be greater the higher the level of union impact, α_u, industry unionism, U, and plant size, PS.

The predicted values for nonunion wages derived from the regression equations actually follow very closely the pattern just described. This is illustrated in Chart I where expected nonunion wages are graphed against the level of industry unionism, assuming mean, high and low values for the union impact variables. The chart is based on equation III.1 and assumes mean or arbitrary values for the control variables, including PS. As can be readily seen, expected nonunion wages increase faster the higher the level of U and the higher the level of α_u. Although not shown on the chart, assuming higher levels of PS has the expected effect of increasing the curvature of the lines. The only area of nonconformity arises because at low levels of U and α_u, nonunion wages decline as U increases while this is not permitted by the specification in (18). However, it should be noted that the declines are modest. For example when low values for the union impact variables are assumed, the expected nonunion wage declines from a value of \$1.76 when U = 35 to a minimum of \$1.70 when U = 60 and rises thereafter to \$1.80 when U = 90.[22]

A similar comparison can be made for increases in plant size. Differentiating (16) with respect to plant size gives: [23]

$$\frac{\partial \alpha_n}{\partial PS} = \alpha_{n_1} U\ \alpha_u \qquad (19)$$

This specification suggests that nonunion wages vary positively with plant size as a result of the effect of spillovers. Ceteris paribus, the increases will be larger the higher the level of U and α_u.

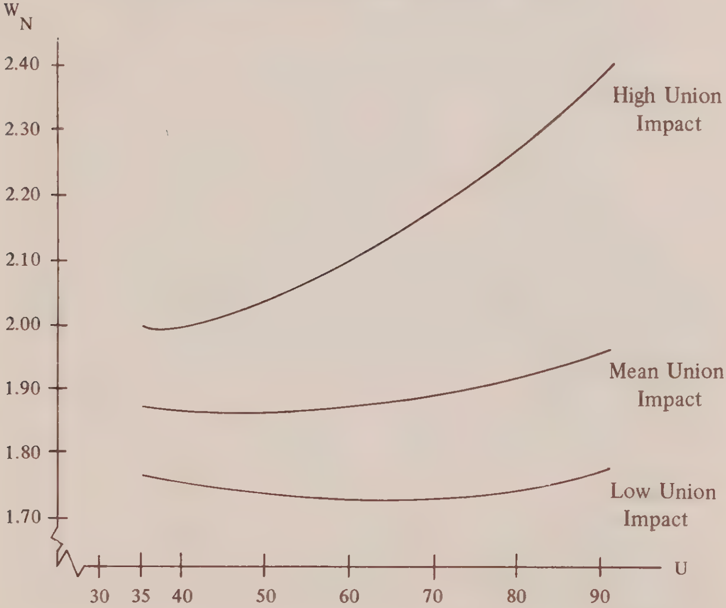
When the predicted changes in nonunion wages due to plant size variation were calculated using the spillover regressors from III.1, it was

[22] The behaviour of expected nonunion wages suggests that perhaps a threshold effect would be involved in a proper specification with all spillover effects being zero until some critical level of U or U·PS is reached.

[23] This assumes that α_u does not depend on PS.

CHART 1

EXPECTED NONUNION WAGE RATES FOR MALE BASIC LABOUR BY EXTENT OF INDUSTRY UNIONISM (DERIVED FROM EQUATION III.1)



found that these last two conditions are satisfied. To illustrate, Table 11 shows expected union and nonunion wages assuming alternate combinations of plant size and industry unionism. Reference groups values for the categorical variables and mean values for all other continuous variables were used in the calculations. Two important patterns emerge. First, nonunion wages increase much less quickly with plant size at low levels of industry unionism than at high levels. In fact, at low levels of industry unionism, increases in nonunion wages associated with plant size increases are quite modest. Second, and really a corollary of the first pattern noted, the level of industry unionism has a greater effect on the wages of large as opposed to small nonunion plants. Both these patterns suggest the presence of spillover effects. Furthermore, the implication is that the positive association between wages and plant size for both union and nonunion employees may in part be attributable to unionism.

TABLE 11

EXPECTED UNION AND NONUNION WAGE RATES FOR FEMALE BASIC LABOUR BY PLANT SIZE AND INDUSTRY UNIONISM BASED ON EQUATIONS II.1 AND III.3

Plant Size	III.1				III.3			
	Union	Nonunion			Union	Nonunion		
		u=35	u=58	u=90		u=35	u=58	u=90
20	2.00	1.83	1.83	1.94	2.00	1.83	1.82	1.93
150	2.14	1.88	1.89	2.08	2.15	1.88	1.89	2.10

It would appear, then, that although the exact specification of (16) is not supported, the regression equation III.1 is consistent in most respects with the existence of spillover effects. Although not discussed, the same conclusion applies to the other equations reported in Table 9.

If the regression equations can be accepted as reasonable approximations for a properly specified model, it is a straightforward matter to make an estimate of the wage of nonunion labour in the absence of spillover effects. This can be done by calculating the expected wage for nonunion labour under conditions where union impact and threats can be presumed to be low. In making these calculations, reference group or mean values are used for the control variables, including plant size. As nonunion wages decline initially with increases in the extent of industry unionism, both a low and the mean value of U are used in the calculations. The previously described set of "low" values for the union impact variables are also used.

The results of these calculations appear in Table 12. All the estimates fall in the range of \$1.70 to \$1.76 per hour and a value of \$1.75 per hour might be taken as the point estimate of the supply price of labour under the conditions set by the control variables.

This estimate can now be used in calculating what has been called the "actual" union-nonunion wage differential. Applying the mean values for the expected union wages in each formulation to the estimate of the supply price of labour yields an inter-establishment differential of nineteen per cent. As was seen in the discussion of the Variant II model, taking into account the distribution of union employees among industries increases the mean value of union wage rates by approximately eight cents per hour. As the regression coefficients of the union impact variables are similar in

TABLE 12

EXPECTED NONUNION WAGES FOR MALE BASIC LABOUR IN THE ABSENCE OF SPILLOVER EFFECTS BASED ON EQUATIONS III.1 AND III.4

	III.1	III.4
	$ per hour	$ per hour
U = 58	1.70	1.70
U = 35	1.76	1.76

the Variant II and III models, the upward adjustments would be virtually identical in the two cases. This means that the mean level of wages for union employees would be approximately $2.16 and that the union-nonunion wage differential in terms of employees would be approximately twenty-three per cent.

Summary

The results reported in this chapter may be briefly summarized as follows. For male basic labour, union establishments, on average, have a wage advantage of about nine per cent relative to comparable nonunion establishments. The size of this differential varies primarily in response to the extent of industry unionism, wages as a proportion of value-added and industry concentration. Plant size and profits appear to have small and perhaps questionable roles in increasing a union's wage gaining ability, whereas there is little evidence to indicate that barriers to entry, as measured by average firm size in an industry, have any impact on wages. When variations in union impact and the distribution of union employees are taken into account, it appears that the bulk of unionized male unskilled labour is employed in industries where the wage advantage of unionism is ten to fifteen per cent. The mean observed differential, in terms of employees, is fourteen per cent. Omitting M as a control variable has the effect of increasing this average differential to twenty-two per cent.

The above estimates would appear to understate the total effect of unionism as there is evidence that unionism has direct effects on the wages

of nonunion establishments. When adjustments are made for these spillover effects, the union-nonunion wage differential in terms of employees, increase from fourteen to twenty-three per cent. But this last estimate may involve a substantial error as the regression results reveal a specification problem in the underlying model.

CHAPTER IV

THE FEMALE BASIC LABOUR RATE

It cannot be certain that unionism affects all groups of workers in the same way. Depending on supply and demand conditions and behaviour patterns, it is quite possible that union effects may vary by sex or skill level. In this chapter the regression results for female basic labour are presented to determine if the union impact is comparable to that found for males. The female basic labour rate will be interpreted as being representative of the wages paid to the twenty-two per cent of all manufacturing production workers in Ontario who happen to be female.[1] In the next chapter variations in union effects by skill level are explored.

Industrial Unionism and Sex Discrimination

Very little attention has been paid to the relationship between union wage effects and sex discrimination at either the theoretical or empirical level. Becker's treatment of discrimination devotes only two pages to the role of unionism and it presupposes craft unions who discriminate on the basis of who is accepted into membership.[2] In manufacturing, industrial unionism is by far the most prevalent form and discrimination by this type of union through exclusion from membership is not practiced. The employer selects who is to be hired and the union admits into membership all those who apply. In this section, possible industrial union effects on female wage rates are discussed in the context of Becker's discrimination model.

Consider, first, a competitive labour market where both males and females are used in production and are paid different wages. This wage differential will reflect both actual productivity differences between males and females, and sex discrimination. The actual productivity differences may relate to differences in the amount and kinds of work performed, expected absenteeism rates, or the expected permanency of attachment with the employer. Sex discrimination by employers may be due to either ignorance or prejudice. Sex discrimination based on ignorance reflects employer misinformation on the actual productivities of males and females,

[1] Dominion Bureau of Statistics, Manufacturing Industries of Canada Summary for Canada, 1968, Catalogue No. 31-203 (Ottawa: Queen's Printer, 1970).

[2] Gary S. Becker, The Economics of Discrimination (Second Ed., Chicago: The University of Chicago Press, 1971), pp. 62-63.

while discrimination based on prejudice probably reflects the belief that males have a prior claim to income as the primary breadwinners with female earnings being viewed as "merely additional income".[3] Whatever the reason, sex discrimination results in employers being willing to pay a higher wage premium for male over female labour than the actual ratio of marginal productivities would justify. Becker has shown that the operation of the labour market will result in a male-female wage differential attributable to sex discrimination that will be larger, the greater the employers' "tastes" for discrimination, the smaller the variance in these tastes, and the greater the ratio of females in the labour force.[4] Employment of females will be concentrated, after making allowances for inter-firm differences in technology, in those areas where employers have the lowest taste for discrimination. But in equilibrium, the actual male-female wage differential due to sex discrimination will be the same throughout the labour market.

An industrial union exercises its power by using the strike threat in negotiations over the wage structure. It must decide how to allocate the wage concessions it can obtain to various groups. In its decisions on the wage structure, the union can either pursue a policy of "equal pay for equal work" or indulge its "taste" for discrimination against females.

It is quite clear that unions at the level of national executives and central federations have endorsed the policy of "equal pay for equal work."[5] If this policy is actually carried out at the bargaining table, then the observed union effect on female wage rates will be larger than that for males.

But, even if a union is sympathetic to the "equal pay" cause, it may feel constrained in pursuing this objective in negotiations. Presumably some unions are aware that a narrowing of the male-female differential will act as an incentive for the employer to substitute male for female employment. Hence a union may be in the position of having to trade off the realization of "equal pay" objectives against a potential loss in female employment. Of course, the size of the loss in female employment will be related to the size of the increase in relative female wages and the ease of substitution between male and female labour.

---

[3] For a fascinating description of attitudes on this question see Renee Geoffroy and Paul Sainte-Marie, Attitude of Union Workers to Women In Industry, Studies of the Royal Commission on the Status of Women in Canada (Ottawa: Information Canada, 1971), pp. 7-28.

[4] Becker, The Economics of Discrimination, pp. 39-54.

[5] See for example: Canadian Labour Congress, "Submission by the Canadian Labour Congress to the Royal Commission on the Status of Women," Ottawa, 1968 (Mimeographed), passim.

On the other hand, the possibility of substitution of male for female labour suggests that, perhaps to a degree, self-interest may lie behind male union member support for the principle of "equal pay for equal work." As one union member put it:

> It is not because they are liberal or soft-hearted that union members favour the idea of wage parity. It is simply that the men are confident that when employers have to pay the same wages to women as to men, they will naturally prefer to hire men.[6]

It is not clear that unions at the local level actually adhere to the principle of "equal pay for equal work." Rank and file union members may in fact share employers' tastes for discrimination.[7] Such a divergence in views within the union hierarchy is illustrated by the following quote:

> At the level of trade-union organizations and federations of labour, we are agreed in principle that female employment is acceptable. It is at the level of the rank-and-file membership that you will encounter this opposition, because they feel their job security threatened and their prejudices will come to the fore.[8]

To the extent that male union members' wishes prevail in decision-making and they hold the view that women's salaries are "merely additional income," there will be a reluctance to allocate wage concessions obtained from the employer toward predominantly or exclusively female occupations. If union members and employers hold the same "tastes" for discrimination, it follows from Becker's analysis that discrimination in

[6] Geoffroy and Sainte-Marie, Attitude of Union Workers to Women in Industry, p. 23. Another quote, in the same vein, goes: "It has been pointed out that the campaign for wage parity between men and women is being fought, above all, by the male workers, and that, as far as they are concerned, the campaign is not motivated by principle but by an instinct of self-preservation. The workers are, in other words, anxious to protect their own jobs, in the belief that employers may prefer to hire women, and women only, to do the work of men for less pay. If, or when that should happen the men will be out of work. On the other hand, the women themselves are reluctant to ask for equal salaries with the men, because they too are afraid in turn that employers may prefer to hire male workers once the wage disparity is removed . . . To sum up, it seems to me that male workers are generally inclined, for reasons of job security, to favour wage parity, while the women, on the other hand, are far less happy about it."

[7] A sizeable gap between the public statements of union leaders and the actual practices of local unions is not necessarily an exceptional event. For another illustration, see Ashenfelter, "Racial Discrimination and Trade Unionism," pp. 440-44.

[8] Geoffroy and Sainte-Marie, Attitude of Union Workers to Women in Industry, p. 13.

unionized markets will be greater than in nonunion markets. The policy selected by a union will generally reflect the median taste for discrimination among its members. But, it is the lower end of the distribution of the taste for discrimination among employers that is critical in determining the size of the male-female differential due to employer discrimination.[9]

However, discrimination by industrial unions may be constrained in a number of ways. First a union may associate discrimination in the form of lower wages for females with a cost in terms of male jobs, if it is aware of the possibility that employers, in the long run, can substitute relatively low priced females for males. Second, if a union's jurisdiction contains a substantial number of females, its policies are less likely to be discriminatory. Even if females constitute a minority within the union, the intensity of their views on this subject may result in their having a disproportionate effect on the union's policy. Finally, a union's ability to discriminate may be hampered by the need to maintain at least nominal adherence to the "equal pay for equal work" principle.[10]

To sum up then, unions, depending upon whether they pursue equal pay objectives or discriminate, may either reduce or increase male-female differentials due to sex discrimination by employers. The correct test of union impact in this area is not a comparison of union wage scales for males and females in one or more job classifications defined as constituting equal work. This approach cannot be conclusive as any observed differential may still be the result of actual productivity differences. Moreover, it would fail to take into account the fact that sex discrimination would probably remain in the absence of unions. The proper question is whether union wage rates, despite the possibility of adverse employment substitution effects, are more or less discriminatory against females than they would be under nonunion conditions. The relative union-nonunion differentials for males and females must be compared.

The regression results for the female basic labour rate are given in the following sections. The presentation follows closely that used in the case of males. In the next section, the Variant I model and coefficients of the control variables are discussed. Subsequent sections deal with the Variant II and Variant III models. Throughout each section, the results are compared to those obtained for males. These comparisons are made easier by

[9] Becker, The Economics of Discrimination, p. 62 and p. 70.

[10] It is doubtful, though, that this principle acts as a severe constraint. For example, nominal adherence might be obtained by simply removing sex labels from jobs, which nonetheless remain segregated on the basis of sex, with no real attempt being made to alter the wage structure.

the fact that the sample of establishments reporting female basic labour rates is identical to that used in the analysis of basic male rates.

The Variant I Model and Control Variables

The regression equation for the female basic wage rate using the Variant I model is as follows:

$$\text{(IV.1)}\quad W = .804 + \underset{(2.6)}{.061\, R_1} + \underset{(4.5)}{.116\, R_2} - \underset{(2.1)}{.055\, R_3} - \underset{(0.0)}{.001\, R_4} + \underset{(14.0)}{.0892\, PS}$$

$$- \underset{(0.6)}{.011\, C} - \underset{(0.5)}{.0079\, D} + \underset{(16.9)}{.00539\, M} + \underset{(4.7)}{.000806\, E} + \underset{(9.4)}{.135\, CA}$$

$$R^2 = .220$$
$$N = 3422$$

The coefficients of the control variables indicate variations in wage rates that are similar to those found for males. Regarding plant location, Region 1 (Central Ontario and Niagara) and Region 2 (Lake Erie and Lake St. Clair) are again identified as high wage areas. However, the wage advantages of these areas are not as large as for males, being only six and twelve cents respectively. As in the case of males, the suppressed category (Eastern Ontario and Lake Ontario) and Region 3 (Midwestern Ontario and Georgian Bay) appear as low wage areas. A difference does arise in Region 4 (Northeastern and Northwestern Ontario). This region appears as a high wage area for males but a low wage area for females. This could arise because in this part of the province the relatively high paying resource industries provide an alternative source of employment for males but not for females. The overall pattern of regional wage differentials was found to be basically the same in the Variant II and III regression equations.

Neither a city location nor work within a durable good manufacturing industry has a marked effect on female wage rates. The coefficients of the corresponding dummy variables are generally small, negative, and statistically insignificant. In some formulations the coefficients of the durable good dummy variable are positive and significant. But in these cases the implied wage effects are still small, being in the range of three to four cents per hour.

As occurred in the regressions for the basic male rate, the ratio of males has a large positive influence on female wages. However, the effect is somewhat smaller. In the male equations, the coefficient of M varies in the range of .0088 to .0069 while the range for the female equations is .0054

to .0040. Again the coefficient displays considerable stability as variables delineating union impact are added to the equations.

Finally, for both male and female basic labour, the plant size coefficients in the Variant I models are about the same. In each case, plant size exerts a strong positive effect on wage rate levels.

Of primary importance for this study is the overall union-nonunion wage differential. The Variant I model indicates that union establishments pay 13.5 cents per hour more than comparable nonunion establishments. Although this is some 4 cents per hour less than the comparable differential for males, the low wages of females result in the relative advantage being almost identical for both sexes. Using the standard reference group yields an expected nonunion female wage rate of $1.69 per hour ($1.97 for males) and a relative differential of 8.0 per cent (9.0 per cent for males).

The Variant II Model – The Role of Union Impact Variables

The regression results for various formulations of the Variant II model are presented in Table 13. Two general characteristics of these results are notable. First, the variables accounting for variations in union impact have the expected signs and are in most cases statistically significant. Second, there is often a marked similarity between the coefficients of the comparable male and female regression equations.

The union impact variables U, W/VA, and CR play prominent roles in the case of the female basic labour equations. A comparison between IV.2 and the corresponding male equation indicates that the coefficients of U and CR are nearly the same for both sexes. However, the coefficient of W/VA in the female equation is only about half its size in the male equation. Another important difference is that when PS is added as a union impact variable (see equations IV.3 and IV.4) its coefficient is large and statistically significant while the coefficient of the plant size control variable is cut in half. In contrast to the situation for males, increases in plant size appear to have a much larger effect on female wages in organized as opposed to unorganized plants. However, for both sexes, there appears to be little evidence supporting interaction effects between concentration and plant size.

Other union impact variables proved to have only limited roles. Permitting interactions between CR and W/VA on the one hand, and U on the other, adds little to the explanatory power of the equations. Moreover, the interaction coefficients are generally small and not statistically significant.

TABLE 13
VARIANT II:
REGRESSION RESULTS FOR FEMALE BASIC LABOUR – EQUATIONS IV.2 TO IV.8

VARIABLES	IV.2	IV.3	IV.4	IV.5	IV.6	IV.7	IV.8
PS	.0717 (11.0)	.0363 (3.4)	.0363 (3.4)	.0720 (11.4)	.0363 (3.4)	.0748 (11.6)	.0722 (11.1)
CA	.130 (2.6)	−.0654 (1.0)	−.0274 (0.4)	−.0466 (0.4)	−.210 (1.4)	.0464 (1.5)	.105 (1.1)
U·CA(10^{-2})	.348 (6.4)	.315 (5.7)	.314 (5.7)	.618 (3.1)	.608 (3.1)	.215	(1.7)
CR·CA(10^{-2})	.415 (3.6)	.404 (3.5)	.0358 (0.1)	.744 (1.9)	.435 (0.9)		.402 (3.5)
W/VA·CA(10^{-2})	−.631 (6.5)	−.644 (6.6)	−.645 (6.6)	−.223 (0.7)	−.222 (0.7)		−.561 (5.6)
PS·CA		.0558 (4.2)	.0474 (3.0)		.0449 (2.8)		
PS·CR·CA(10^{-2})			.0494 (1.0)		.108 (1.3)		
CR·U·CA(10^{-2})				−.00438 (0.9)	−.00703 (1.4)		
W/CA·U·CA(10^{-2})				−.00630 (1.4)	−.00645 (1.4)		
R·CA(10^{-2})						−1.6 (4.1)	−.0609 (0.1)
R·U·CA(10^{-2})						.0436 (9.6)	.0144 (1.3)
R^2	.254	.257	.258	.254	.258	.245	.256

The profit variables, while significant when appearing alone as the only union impact determinant, appear to have only limited effects when other variables are included.

In summary then, the roles played by the union impact variables appear to be quite similar for both male and female unskilled labour with W/VA, CR, and U having the largest effects. But in the case of females, PS also appears as an important determinant of union impact.

The latter result implies that the union-nonunion wage differential obtained from the Variant I model is biased downwards. In that regression, the plant size control variable is actually picking up some of the variation in wage rate levels that ought to be associated with union status. The coefficient of the CA dummy variable is accordingly too small. An unbiased estimate can be obtained by using the Variant II regression coefficients and the mean values for the union impact variables to calculate expected union and nonunion wage rates. Using the standard reference group and the coefficients from equation IV.3 results in an expected nonunion wage of 1.643 and an expected union wage of 1.788. The differential of 14.5 cents is one cent higher than obtained from the Variant I model.

In order to determine the union-nonunion wage differentia in terms of employees rather than establishments, the same procedure as used in Chapter III was followed. For each industry, the coefficients of the union impact variables from IV.3 were used to calculate a union-nonunion wage differential.[11] These differentials were then weighted by an estimate[12] of female union membership in the industry. The results of these calculations appear as Table 14.

TABLE 14

PERCENTAGE DISTRIBUTION OF FEMALE UNION EMPLOYEES BY THE SIZE OF OBSERVED UNION-NONUNION WAGE DIFFERENTIALS FOR FEMALE BASIC LABOUR

Differential (cents per hour)	Female Union Employees (per cent)
Less than – 9	2.8
– 9 to – 1	3.6
0 – 9	22.4
10 – 19	16.9
20 – 29	33.7
30 – 39	15.3
40 – 49	2.3
50 – 59	0.8
60 – 69	2.1
70 – 79	0
Weighted Average = 20.0	

[11] In making the calculations, the natural log of the mean plant size for union establishments in an industry was used rather than the correct variable, the mean of the natural log of plant size for all union establishments. This imparts a mild upward bias to the estimates – the degree depending upon the extent of inequality in union plant size for the industry. Even when extreme assumptions are made concerning inequality in plant size, the resulting bias is less than 3 cents per hour and undoubtedly averages much less than this in the actual sample of industries.

[12] The ratio of female production workers times total non-office employment in union establishments as reported on the Canada Department of Labour survey tape.

Again it would appear that union effects vary considerably. In fact, although somewhat similar, the distribution of union effects for females appears less concentrated about the mean value than the same distribution for males. In the latter case, almost half the male employees were found to have union effects that were in a range of plus or minus five cents per hour about the mean value. Whereas in the female distribution, union effects throughout the range of 0 to 40 cents per hour are quite common. On the other hand, the distribution for female employees shows fewer extreme values. Less than 12 per cent of female union members are subject to union effects outside the range indicated above.

Taking account of the distribution of female union members has the effect of raising the average differential from 14.5 to 20.0 cents per hour. In relative terms the differential increases modestly to 12.2 per cent. In both cases, these differentials turn out to be remarkably close to those found for males, although somewhat smaller. The clear implication appears to be that existing sex discrimination in the labour market can only to a small degree be attributed to a different impact of unionism on the wage scales of men and women.

The finding of no large difference in union impact by sex is in line with the results reported in the three available studies on the United States experience. They all found that the male and female differentials were quite similar, although in each case the female differential turned out to be somewhat larger.

The Variant III Model – Spillover Effects

The similarity of results for males and females extends to the Variant III regressions. Table 15 presents the results for three formulations of the model in which it is assumed that plant size is one of the union impact variables and that, alternatively, threat relations are governed by:

(1) $\alpha_n = \alpha_{n_0} + \alpha_{n_1}\, U{\cdot}PS\, \alpha_u$; equation IV.9

(2) $\alpha_n = \alpha_{n_0} + \alpha_{n_2}\, U\, \alpha_u$; equation IV.10

(3) $\alpha_n = \alpha_{n_0} + (\alpha_{n_1}\, U{\cdot}PS + \alpha_{n_2}\, U)\, \alpha_u$; euqation IV.11

The corresponding Variant II model is also presented for comparative purposes.

Considering first equations IV.9 and IV.10, all the spillover effect coefficients for which we have an a priori expectation have the proper sign,

TABLE 15

VARIANT III: REGRESSION RESULTS FOR FEMALE BASIC LABOUR – EQUATIONS IV.2, IV.9, IV.10 AND IV.11

VARIABLES	IV.2	IV.9	IV.10	IV.11
PS	.0363 (3.4)	.0086 (0.4)	.0128 (0.4)	.0435 (0.8)
CA	-.0654 (1.0)	-.2441 (2.6)	-.2277 (1.8)	-.1158 (0.5)
$U{\cdot}CA(10^{-2})$	.315 (5.7)	.314 (5.8)	.1314 (5.8)	.314 (5.8)
$CR{\cdot}CA(10^{-2})$	.404 (3.5)	.414 (3.6)	.413 (3.6)	.414 (3.6)
$W/VA{\cdot}CA(10^{-2})$	-.644 (6.6)	-.685 (7.1)	-.686 (7.0)	-.683 (7.1)
$PS{\cdot}CA$	.0558 (4.2)	.0832 (3.3)	.0791 (2.6)	.0484 (0.9)
$U{\cdot}PS(1\text{-}CA)\ (10^{-3})$		-1.417 (2.2)	.374 (6.7)	- 3.308 (1.4)
$CR{\cdot}U{\cdot}PS(1\text{-}CA)\ (10^{-3})$		.0313 (4.6)		.00965 (0.4)
$W/VA{\cdot}U{\cdot}PS(1\text{-}CA)\ (10^{-3})$		.0283 (5.6)		-.0275 (1.3)
$U^2{\cdot}PS(1\text{-}CA)\ (10^{-3})$		.0148 (2.7)		.00723 (0.3)
$U{\cdot}PS^2(1\text{-}CA\ (10^{-3})$		.2096 (2.6)		.4410 (3.8)
$U(1\text{-}CA)\ (10^{-3})$			- 3.777 (1.4)	3.332 (0.4)
$U^2(1\text{-}CA)\ (10^{-3})$			.0604 (2.9)	.0334 (0.4)
$U{\cdot}CR\ (1\text{-}CA)\ (10^{-3})$			.1165 (4.6)	.0774 (0.8)
$W/VA(1\text{-}CA)\ (10^{-3})$			-.1077 (5.6)	-.00137 (0.0)
R^2	.257	.280	.278	.282

and in all but one instance are statistically significant. Estimates of α_{n_1} based on equation IV.9 fall in the range of .0025 to .0076. This is quite close to the estimates derived from the corresponding equation involving the basic male rate. The estimates of the structural parameter α_{n_2}, based on equation IV.10, also appear quite reasonable. The four estimates range from .005 to .028. In terms of the goodness of fit, there is little to choose between the two specifications. Both sets of spillover effect regressors result in a statistically significant reduction in the residual sum of squares over the Variant II model.

Turning to equation IV.11 it is evident that substantial multicollinearity among the spillover effect regressors is present. As a result, large standard errors reduce the precision of the parameter estimates. The implicit values of α_{n_1} and α_{n_2} vary greatly, although generally they do have the proper sign.

All the formulations, however, suffer from the same apparent misspecification as appeared in the male equations. The negative coefficients for U·PS(1-CA) in equations IV.9 and IV.11 and U(1-CA) in IV.10 imply negative or small values for α_{u_0}. This, in conjuction with normal values for the other variables, suggests negative spillover effects – a result inconsistent with the underlying model. Clearly, little reliance can be placed on the particular specification of the model used in the regressions.

Nonetheless, the predicted values of nonunion wages do follow a pattern consistent with the existence of spillover effects. In Table 16 predicted values for union and nonunion wages under conditions that favour a high, average, and low union impact are presented. Under conditions that favour either a large union impact and spillover effect or a small union impact and spillover effect, union and nonunion wages appear to be quite similar. On the other hand, where union effects are appreciable but spillover effects moderate, union and nonunion wages diverge by a substantial amount.

In Chapter III it was noted that the model implies that if spillover effects exist, an increase in industry unionism, U, should result in an increase in nonunion wages and that the increase should be larger the higher the level of union impact, α_u; plant size, PS; and U. Chart 2 displays the behaviour of nonunion female basic labour wages as U increases derived from equations IV.9 and IV.10. Again, it was found that nonunion wages do follow the expected pattern with the exception that at low levels of U and α_u increases in U actually lead to small decreases in nonunion wages.

Further support for the presence of spillover effects is provided by the influence of plant size and industry unionism effects on nonunion wages.

TABLE 16

EXPECTED UNION AND NONUNION WAGES FOR FEMALE BASIC LABOUR BASED ON EQUATIONS IV.9, IV.10 AND IV.11

	IV.9		IV.10		IV.11	
Union Impact*	Union	Nonunion	Union	Nonunion	Union	Nonunion
High	2.13	2.25	2.13	2.10	2.14	2.16
Mean	1.77	1.59	1.77	1.61	1.77	1.58
Low	1.47	1.52	1.47	1.49	1.48	1.50

*The values of the variables used in the calculations correspond to those used and reported in Chapter III.

This is best illustrated through an examination of expected nonunion wages under varying conditions. From Table 17 it can be seen that increases in industry unionism have a larger effect on the wages of large as opposed to small nonunion plants and that increases in plant size only have particularly large effects on nonunion establishments located in well-organized industries. Both these patterns suggest the presence of spillover effects.

TABLE 17

EXCEPTED UNION AND NONUNION WAGE RATES FOR FEMALE BASIC LABOUR BY PLANT SIZE AND INDUSTRY UNIONISM BASED ON EQUATIONS IV.9, IV.10, AND IV.11

	IV.9				IV.10				IV.11			
Plant Size (Employees)	Union	Nonunion			Union	Nonunion			Union	Nonunion		
		U=35	U=58	U=90		U=35	U=58	U=90		U=35	U=58	U=90
20	$1.67	$1.57	$1.57	$1.67	$1.67	$1.59	$1.54	$1.63	$1.67	$1.57	$1.58	$1.68
150	[illegible]	[illegible]	[illegible]	[illegible]	[illegible]	[illegible]	$1.57	$1.80	$1.95	$1.62	$1.63	$1.76

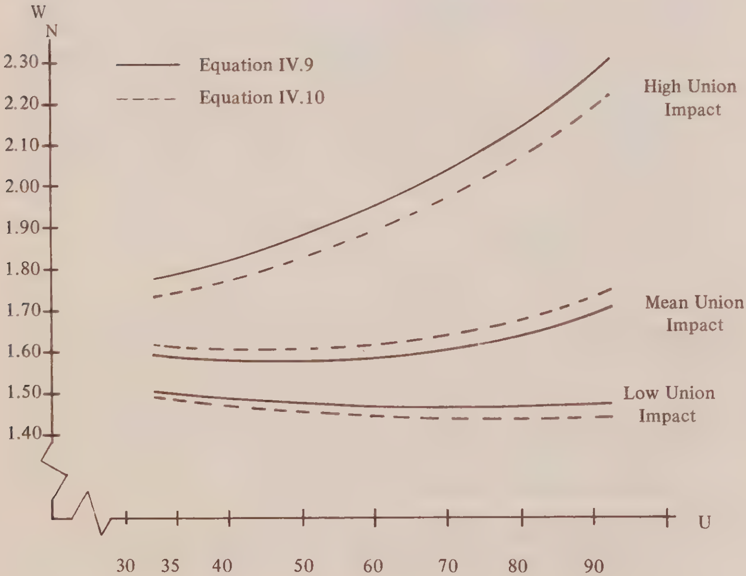
Following the procedure used in Chapter III, the Variant III regression results can be used to calculate an estimate of the actual union – nonunion wage differential. Using the standard reference group and assuming conditions that favour a small spillover effect, including a small

plant size, yields an expected nonunion wage of \$1.49 to \$1.52 depending upon the equation used. The comparable union rate, evaluated at mean levels for the union impact variables is \$1.77. Taking account of the distribution of female union members raises the differential to 34 cents. In relative terms, the actual union-nonunion wage differential in terms of employees is 23 per cent – the same as the male differential.

Again, this point estimate of the actual union-nonunion wage differential must be regarded as highly tentative. Not only is there an apparent specification error in the underlying model but also the union impact and spillover effect regressors may be capturing influences other than unionism on wage rates. To the extent that this is true, these estimates tend to overstate the influence of unionism.

CHART 2

EXPECTED NONUNION FEMALE BASIC LABOUR WAGE RATES BY EXTENT OF INDUSTRY UNIONISM (DERIVED FROM EQUATIONS IV.9 AND IV.10)



Implications of the Results

The apparent similarity of union impact on male and female unskilled wage rates demonstrated throughout this chapter does not necessarily mean that unionism has had no effect in increasing or decreasing the inequality between male and female wages. The latter effect depends on the extent to which male and female workers are organized as well as the proportionate effect of unionism on the wages of male and female workers.[13] However, the available evidence suggests that at least within manufacturing about the same proportion of males and females have elected to join unions. Females comprise 19 per cent of all union members and about 23 per cent of all production workers in Canadian manufacturing.[14] Also the similarity between the spillover effects for males and females strongly suggests that unionism has affected the wages of nonunion workers of both sexes in roughly the same proportion. Therefore, the implication of the results of this study is that unionism has not materially affected the degree of inequality between male and female wages in manufacturing.

The demonstrated similarity between the union-nonunion wage differentials for males and females provides rather strong evidence that unions within manufacturing have not exercised their power either to discriminate against females or raise female wages so as to bring about "equal pay for equal work." But this result does not imply that unions are unsympathetic toward the latter cause. First, the results of this study are for the year 1969. Judging from casual observations, it would appear that the "equal pay" movement has gathered strength among unionists since this date. Second, it was noted that if employers and unionists actually

---

[13] Let the aggregate effect of unionism on the average wage of female labour relative to male labour be defined as,

$$\Delta = [(W_F/W_M)U - (W_F/W_M)C] \; / \; (W_F/W_M)C$$

where $(W_F/W_M)U$ and $(W_F/W_M)C$ are the ratios of female to male wages in the presence and absence of unionism respectively. It can be shown that $\Delta \simeq F \cdot D_F - M \cdot D_M + (T_F - T_M)$

where: D_M and D_F are the observed union-nonunion differentials for males and females relative to nonunion wages, F and M are the proportions of females and males who are union members, and T_F and T_M are the proportionate effects of unionism on the wages of nonunion females and males. See Ashenfelter "Racial Discrimination and Trade Unionism," pp. 436-38. The argument of this paragraph is that for manufacturing industries in Ontario $\Delta \simeq 0$.

[14] See Statistics Canada, Annual Report of the Minister of Industry, Trade and Commerce under the Corporations and Labour Unions Returns Act, 1969, Part II, (Ottawa: Queen's Printer, 1971), p. 67.

share the same “tastes” for discrimination one would expect to find greater wage discrimination in union as opposed to nonunion markets. As the male – female differential is the same or only slightly larger in union markets it would seem safe to conclude that the “tastes” for sex discrimination do not differ greatly between unionists and employers or that unionists are more sympathetic toward equal pay objectives than employers. Finally, it should be kept in mind that industrial unions, lacking control over the employment decision, may in fact be effectively constrained in the pursuit of equal pay objectives by possible adverse substitution effects. Even for a union which seeks to combat sex discrimination, it would make little sense to demand equal wages for males and females if the expected effect is a substantial reduction in female employment.

CHAPTER V

UNION-NONUNION WAGE DIFFERENTIALS BY SKILL LEVEL

The current state of knowledge concerning the impact of unionism on skill differentials can be illustrated by a quote from a recent paper by Lewis:

> Economists for some time have agreed to disagree about the effect of unionism on the relative wage structure by occupation or skill level. I regarded the evidence in my book on this question as far too scanty to settle this question or even sufficient for me to estimate provisionally the direction (widening or narrowing) of the effect. Perhaps I was over cautious, for others have read the evidence as supporting the contention that unionism has widened skill differentials. However, I have recently gone over the evidence again and have come out in the same place. The question is unsettled.[1]

Alternative Predictions

The differing views concerning the expected effect of unionism arise due to the particular assumptions made concerning union behaviour. Rosen[2] analyzes the effects of unionism on inter-occupational wage differentials with the aid of a formal model in which it is assumed that the main objective of a union is to maximize income of employed members over and above what they could earn in alternate pursuits. It follows from this that unions would seek to take advantage of differences in demand conditions between groups of labour. The desired relative union-nonunion wage differential for one group in relation to the same differential for another group is shown to depend on relative labour demand elasticities adjusted for cross demand elasticities between groups of labour. As it is widely held that skilled workers are subject to a more inelastic demand than other production workers, the implied hypothesis is that the relative union-nonunion wage differential will be greater for skilled workers. Rosen

---

[1] Lewis, "Unionism and Relative Wages Revisited," (paper presented at the University of Chicago Workshop Group in Labor Economics and Industrial Relations, May 29, 1967), p.28.

[2] Rosen, "Unionism and the Occupational Wage Structure," p.269-73.

goes on to establish that this differential union impact by skill level is independent of the overall size of union effects as influenced by such factors as product market monopoly and the extent of industry unionism and also independent of the form of union organization. The same conclusions hold even if the skilled and less skilled workers of a single enterprise are organized into separate unions rather than under one union.

This predicted effect of unions is at odds with the widely held view among institutional labour economists that unions are egalitarian institutions which attempt to reduce or eliminate various types of differentials, including those based on skill. The common conception is that industrial unions have a tendency to negotiate equal cents per hour increases for all workers in the bargaining unit, with special skill adjustments being agreed to only reluctantly.[3] The result, of course, is a compression of relative skill differentials.

If equal mark-ups is the predominant pattern, we would expect to find the same absolute union-nonunion wage differential for all occupations. Moreover, the determinants of union impact would have approximately the same effects on union-nonunion wage differentials across all occupations.

Yet another prediction arises from what has been called the "modified neoclassical" view of the labour market.[4] According to adherents of this view, labour markets behave competitively under conditions of excess demand but institutional factors predominate where markets are characterized by excess supply. It is then noted that skilled workers are in fairly inelastic supply to particular markets because there are obvious difficulties in enlarging their numbers over the short run. On the other hand, less skilled workers tend to be in highly elastic supply at institutionally

[3] For example, Peitchinis observes that, "Industrial unions, where membership is predominately unskilled and semi-skilled, have consistently followed a wage policy whose effect on differentials has been to narrow them. Generally, they have insisted on equal distribution of the total increment in wages amongst all workers regardless of skill." See S.G. Peitchinis, The Economics of Labor (Toronto: McGraw-Hill, 1965), pp. 335-36.

[4] See George E. Johnson, "Wage Theory and Inter-Regional Variation," Industrial Relations, Vol. 6, No. 3 (May 1967), pp. 321-338; Pradeep Kumar, "Differentials in Wage Rates of Unskilled Labour in the Manufacturing Industries," Industrial and Labor Relations Review, (forthcoming); G.H. Hilderbrand and G.E. Delehanty, "Wage Levels and Differentials" in Prosperity and Unemployment, eds. R.A. Gordon and M.S. Gordon, (New York: John Wiley, 1966), pp. 265-301; S. Behman, "The Wage Determination Process in U.S. Manufacturing," Quarterly Journal of Economics, Vol. LXXXI, No. 1 (February 1968), pp. 117-42; R.L. Raimon, "The Indeterminateness of Wages of Semiskilled Workers." Industrial and Labor Relations Review, Vol. 6, No. 2 (January 1953), pp. 180-94.

determined wage rates, as long as general excess supply conditions for this type of labour prevail. This, it is argued, is the normal state of affairs due to downward wage rigidity. As a consequence the wage structure, in response to demand shifts, serves as an allocative mechanism for skilled occupations but not for the unskilled. The major union impact is to push up the wages of unskilled workers in line with skilled rates as a result of the pursuit of "equitable" internal wage structures.

To the extent that this type of dual mechanism operates it would be expected that the overall union-nonunion wage differential would be much less for skilled workers. In addition, it would be expected that the variation in union skilled rates would be less than in the case of the unskilled while their sensitivity to demand shifts would be greater.

In order to put these various predictions to the test, the models used in Chapters III and IV were run on a set of narrowly defined occupations. These occupations fall into three basic categories: skilled, semi-skilled and unskilled. Not all establishments in manufacturing report rates for each of the occupations. Consequently, the sample of establishments (and industries within manufacturing) varies between the occupations. This exercise should also serve to determine the extent to which the regression results obtained for basic male and female labour are stable over more narrowly defined occupational categories.

Variant I and II Model Results

From the Variant I regressions, the coefficient of the dummy variable, CA, gives the size of the union-nonunion wage differential in absolute terms. As can be seen from Table 18, there is a clear inverse relation between the size of this regression coefficient and skill level. The differential for the three unskilled occupations ranges between 14 to 15 cents per hour or only slightly below the differential found for the basic male rate. The differentials for the three semi-skilled occupations display more variability ranging from 8 to 16 cents per hour. But the differentials for the skilled occupations are clearly small. For three out of the four occupations, the differentials are less than 5 cents in absolute terms and not statistically different from zero. The other skilled occupation shows a modest differential of 11 cents.[5]

[5] No formal hypothesis test of the differences in the coefficients by skill level was carried out. Given the costs involved in running the data on a pooled basis, such a test was judged to be impractical.

TABLE 18

VARIANT I AND II: REGRESSION RESULTS FOR SELECTED OCCUPATIONS

Occupation	Variant I		Variant II							Number of Observations
	CA	R^2	PS	CA	CR·CA (10^{-2})	W/VA·CA (10^{-2})	U·CA (10^{-2})	PS·CA	R^2	
Skilled:										
Electrician	-.046 (1.0)	.257	.155 (4.1)	.109 (0.5)	-.0813 (0.3)	-.908 (4.1)	.357 (2.9)	-.0119 (0.3)	.281	819
Machinist	.0048 (0.1)	.226	.131 (4.1)	.190 (0.9)	-.693 (2.6)	-1.232 (4.8)	.433 (3.1)	.0124 (0.3)	.264	671
Mechanic (Machine Repair)	.108 (3.1)	.204	.063 (2.3)	-.197 (1.2)	-.0105 (0.0)	-1.060 (5.5)	.532 (4.3)	.0792 (2.4)	.249	1137
Welder	.026 (0.5)	.418	.150 (4.8)	-.306 (1.5)	.317 (1.2)	-.728 (2.8)	.612 (4.3)	.0409 (1.1)	.465	556
Semi-Skilled:										
Shipper	.0831 (4.0)	.214	.0555 (3.6)	.134 (1.4)	.0309 (0.2)	-1.027 (7.3)	.325 (4.1)	.032 (1.7)	.243	2191
Industrial Truck Operator	.108 (3.9)	.303	.176 (8.0)	.655 (5.2)	.454 (2.7)	-1.106 (7.1)	.196 (2.3)	-.755 (3.0)	.355	1053
Truck Driver, Heavy Truck	.160 (4.7)	.256	.109 (4.5)	.579 (3.8)	.577 (2.5)	-1.205 (5.7)	.140 (1.1)	-.032 (1.1)	305	733
Unskilled:										
Cleaner	.149 (6.0)	.279	.114 (5.7)	.212 (1.8)	.414 (2.3)	-1.031 (6.7)	.281 (3.4)	.026 (1.1)	.322	1406
Labourer	.152 (6.1)	.320	.107 (5.4)	.335 (3.0)	.089 (0.5)	-1.172 (7.7)	.328 (3.9)	.014 (0.6)	.365	1123
Labourer (Non-production)	.136 (3.9)	.417	.143 (5.6)	.284 (1.9)	.421 (2.0)	-.962 (5.1)	.357 (3.3)	-.014 (0.5)	.465	674

These results, suggesting that unionism has narrowed both relative and absolute skill differentials, are inconsistent with the type of union maximizing behaviour hypothesized by Rosen. But the analysis offers no clear basis for choosing between the other two hypotheses discussed. As the regressions suggest a small but probably positive union-nonunion wage differential for the skilled occupations the results appear to lie between what was expected from the "equal mark-up" or the "modified neo-classical" viewpoints. To explore the question further it is necessary to examine the coefficients of the union impact variables in the Variant II model. It will be recalled that the former hypothesis suggests a similarity in the coefficients across all skill levels while the latter predicts relatively small or negligible coefficients for the union impact variables and a greater sensitivity to demand factors for the skilled occupations.

First consider the effect on wages of plant size indicated by the regression coefficients reported in Table 18. Although the coefficients of the plant size control variables are larger than those found in the basic male and female regressions, there is little evidence that, within the group of specific occupations, plant size effects vary inversely with skill level. If anything, the effects are marginally larger for the skilled occupations. The union impact variable coefficients also fail to show such a negative relation. In fact, no relation between plant size and union impact is apparent. The coefficients are typically small, statistically insignificant and inconsistent in sign. The one unusually large positive coefficient and the unusually large negative coefficient balance off correspondingly low and high coefficients for the plant size control variables. It seems that, at all skill levels, plant size has the same large positive effect on the wages of both union and nonunion establishments.

In contrast, the effect of the concentration union impact variable does vary by skill level. For three out of the four skilled occupations the relevant coefficient is negative rather than positive as expected. In the skilled occupation where the effect is positive, the coefficient is not statistically significant. For the six semi-skilled and unskilled jobs, the effect is positive in five cases. In the four cases where the coefficient is both positive and significant, its size is approximately the same as found in the basic male and female regressions. The implication of these findings is that concentration is a more significant determinant of unskilled as opposed to skilled wages.

Turning now to wages as a proportion of value-added, it is found that the coefficients display a remarkably consistent pattern across all occupations. All are negative, large in absolute value, statistically significant and within the narrow range of – .00728 to – .01232. There is no

discernible tendency for the coefficients to vary with skill level.

A peculiar pattern is found in the industry unionism coefficients. Although all are positive, and, in all but one case, statistically significant, the size of the effects are smallest for the semi-skilled occupations and largest for the skilled occupations. Not too much emphasis should be placed on this unexpected result as the coefficients for both the skilled and unskilled occupations are roughly in line with those obtained from the basic male and female regressions. If there is a positive association with skill level, it probably is very weak.

In total there is little evidence that the coefficients of the union impact variables are smaller for the skilled occupation. The results, therefore, appear rather anomalous. The overall small size of the differentials in absolute terms suggests that perhaps the "modified neoclassical" hypothesis might in large part explain the limited union effect on skilled wage rates. But the similarity in the coefficients of the union impact variables across all skill levels suggests that the equal mark-up mechanism is operating.

One explanation of this apparent contradiction may be connected with the role played by the ratio of males control variable. It will be recalled that in the discussion of the male basic labour regressions, it was remarked that the effect of this variable is surprisingly large given its rationale as a crude index of working conditions. This large effect raises the possibility that a high ratio of females may in large part be more the result rather than the cause of low wages. If employers with a low wage structure, whatever the cause, are forced to hire relatively large numbers of females, the ratio of males will not be independent of the wage rate level. The model, as interpreted above, is misspecified.

It is noted in Chapter III that omitting M has the effect of increasing the size of the coefficients of the union impact variables. It can be surmised that omitting this variable would have similar effects on the other regression reported in this study.

With this in mind it is important to note that while the effects of M in the semi-skilled and unskilled equations were about the same as observed for the basic male and female rates, the effects for the skilled occupations were quite small in three out of four cases. If the proper specification required that M be omitted, it is likely that the resulting regressions would show relatively small coefficients for the union impact variables in the skilled occupations. The apparent contradiction, mentioned above, would disappear. Unfortunately, the regressions for the skilled occupations were not run with M omitted so that this explanation cannot be put to a firm test.

The applicability of the "modified neoclassical" view can be considered further through an examination of the coefficients of other control variables. To the extent that the modified neoclassical mechanism operates it is expected the coefficients reflecting demand influences would be larger for the skilled occupations. The regression results from the Variant I model indicate that this is, in fact, true, although the differences by skill level are small.[6] The coefficients of the industry growth variable for the four skilled occupations average .001702 while the same averages for the semi-skilled and unskilled occupations are .001256 and .001375. The skilled wage rates also display more variation by region but again the differences by occupation do not appear large. This is illustrated in the following tabulation where average coefficients by skill level are given for the four regional dummy variables.

Skill \ Region	R_1	R_2	R_3	R_4
Skilled	.348	.389	–.009	.370
Semi-Skilled	.275	.295	.062	.269
Unskilled	.173	.230	–.064	.115

To summarize the results so far, unionism within manufacturing appears clearly to be associated with a narrowing of relative skill wage differentials. The extent to which absolute skill wage differentials have been reduced is less clear as the results lie between what may have been expected from the "equal mark-up" and "modified neoclassical" views. The mean differential, although positive, is probably smaller than that observed for the unskilled occupations. But, depending upon the role assigned to the ratio of males variable, the variation in union rates may be viewed as more or less constant across skill levels. Skilled wages appear only slightly more sensitive to demand conditions than the unskilled rates. No definite choice between the two hypotheses appears possible.

[6] As the control variable coefficients do not change radically between different formulations, the conclusions of this paragraph are not changed with the use of either the Variant II or III regression results.

The Variant III Model – Spillover Effects

In order to consider possible union spillover effects, the Variant III regressions were also run on the selected occupations. But, as in the case of the basic male and female categories, a specification problem was found to be present. Although the coefficients generally have the expected signs, the overall contribution of the spillover regressors at mean values often suggests unacceptable negative spillover effects. The individual estimates of the structural parameters also display considerable variation. Again the selected specifications of the spillover relationship were found to be wanting. It is, therefore, impossible to calculate the actual union-nonunion wage differentials based on the estimates of the structural parameters and to make comparisons by skill level. It is possible, however, to determine if the predicted variations of nonunion wages do follow patterns consistent with the presence of spillover effects.

As in the prior two chapters, the questions that must be asked are:

(1) Does the addition of the spillover effect regressors lead to a statistically significant reduction in the residual sum of squares?
(2) Do nonunion wages increase with the extent of industry unionism, U?
(3) Are plant size effects on nonunion wages larger in highly organized industries as opposed to poorly organized industries?
(4) Are nonunion wages high under other conditions that favour a high union impact?

An affirmative answer to each of these questions is expected from all the formulations of the spillover relationship used in this study.

Table 19 presents some results bearing on these issues derived from two formulations of the Variant III regressions. First, consider the effect of the addition of the spillover regressors. For the skilled occupations, the F-statistics are low and the null-hypothesis for the set of coefficients of the spillover regressors cannot generally be rejected at reasonable confidence levels. As lower skilled occupations are considered, the F-values increase and the null-hypotheses can be safely rejected. The spillover regressors appear to be much more important in explaining variations in nonunion wages for the low skilled occupations.

Nonunion wages should vary positively with the extent of industry unionism as both threats and union impact are larger in well-organized industries. This was generally found to be the case for most occupations except those in the semi-skilled group. The relationship was tested by

TABLE 19

VARIANT III: RESULTS BASED ON REGRESSIONS FOR SELECTED OCCUPATIONS

OCCUPATION	F-Statistic For Contribution of Spillover Regressors To The Explained Sum of Squares		Do Non-union Wages Increase With U?		Are Plant Size Effects Larger For Nonunion Firms at High Levels of U?	
	(1)	(2)	(1)	(2)	(1)	(2)
Skilled:						
Electrician	1.9	2.6*	Yes	Yes	No	No
Machinist	1.5	1.2	Yes	Yes	Yes	Yes
Mechanic (Machine Repair)	1.9	2.2*	Yes	Yes	Yes	No
Welder	2.2	1.6	No	Yes	Yes	No
Semi-Skilled:						
Industrial Truck Operator	4.0**	3.33**	No	No	No	No
Truck Driver, Heavy Truck	2.9*	2.3*	No	Yes	No	No
Shipper	6.9**	4.5**	No	No	Yes	Yes
Unskilled:						
Cleaner	9.1**	8.4**	Yes	Yes	Yes	Yes
Labourer (Production)	15.5**	16.9**	Yes	Yes	Yes	Yes
Labourer (Non-production)	12.8**	12.4**	Yes	Yes	Yes	Yes

calculating the change in nonunion wages associated with increases in U over the range of 35 to 90 per cent.[7] Averaging the wage changes calculated for each occupation in each skill group yields thirteen and seventeen cent increases for the skilled and unskilled respectively. The wage change for the semi-skilled group was found to be negligible. The comparable average turned out to be minus three cents.

Nonunion wages are expected to increase with plant size at a faster rate in highly organized industries due to the increased exposure of large plants in these industries to union organization. This was found to be true for all the unskilled occupations except in one instance where a virtually constant rate of increase was found. However, in a number of cases for the semi-skilled and skilled occupations, no such relationship was found.[8]

For virtually all occupations, nonunion wages were found to be higher in circumstances that favour a large union impact. In all the regressions, nonunion wages vary negatively with wages as a proportion of value-added, and, in sixteen out of the twenty Variant III regressions run, a positive association between concentration and nonunion wages is observed.

To sum up, the evidence based on the selected occupations again suggests spillover effects for the unskilled labour groups. For the higher skilled categories, the evidence favouring the presence of spillover effects is considerably weaker. However, this perhaps is what should be expected if it is true that the union impact on the skilled occupation is generally small.

Summary

The evidence presented in this chapter clearly suggests that unionism within manufacturing has led to a narrowing of relative skill differentials. This is in accord with most other studies of the United States experience but is in direct conflict with Rosen's hypothesis and related empirical evidence [9] which indicate that industrial unions exploit the relatively inelastic demand for skilled labour. Our evidence appears to go further than the results of most other studies in suggesting that even the absolute union-nonunion wage differential is lower for the highly skilled trades. Only Ashenfelter's[10] study could be interpreted as going this far.

[7] In the calculations mean values for CR and W/VA were used along with a high (5) and, alternatively, a low value (3) for the plant size variables.

[8] The conclusions of this paragraph are based on the calculations described above.

[9] Rosen, "Unionism and the Occupational Wage Structure", pp. 269-86.

[10] Ashenfelter, "Racial Discrimination and Trade Unionism", pp. 435-64.

Our findings lie somewhere between what might be expected from an "equal mark-up" view of union behaviour, where all groups in the bargaining unit receive the same cents per hour increases, and the "modified neoclassical" view which maintains that skilled trades have their wages determined in a competitive labour market environment. The mean absolute wage differential, although positive, is smaller for the skilled trades than for the unskilled and semi-skilled occupations. In addition, depending upon the role assigned to the ratio of males variable, the significance of the union impact variables may be viewed as more or less constant across skill levels. But, in any case, the union impact variables do appear to play at least some role in determining skilled worker wage rate levels. Two other considerations are that the wages of skilled workers appear only slightly more sensitive to labour market conditions as indexed by demand variables than the unskilled worker rates, and that the evidence favouring the presence of spillover effects is considerably weaker for the skilled trades. All in all, neither of the two extreme views fits the results particularly well.

What is clear is that union impact is negatively related to skill level and that this fact should be taken into account when a global differential for all production workers is being considered.

CHAPTER VI

THE UNION IMPACT ON WHITE-COLLAR WAGES

Direct and Spillover Effects

Unionism can affect the wages of white-collar workers in two ways. First, there is the direct wage effect associated with the unionization of these workers. Second, the unionization of an employer's blue-collar workers may alter the wage decisions he makes for his unorganized workers. Both these possibilities must be considered in assessing union impact in the white-collar field.

The direct effect through organization of white-collar workers appears limited at the present ime. Despite a recent rapid expansion of white-collar unionization among government workers and certain professions, inroads in the manufacturing sector have been few. In the wage rate survey used in this study, less than five per cent of the manufacturing establishments in Ontario reported having a collective agreement covering any office employees. It is also not clear what power can be exercised by organized white-collar workers in manufacturing. Their position may be weakened by a disinclination to strike as white-collar workers typically identify more closely than others with management attitudes and objectives. Moreover, their ability to inflict economic harm on the employer in short-term work stoppages may be restricted by the availability of supervisors who can carry out essential production and the fact that many tasks performed by white-collar workers can be deferred.

Even though weak in their own right, white-collar unions may gather strength from their association with blue-collar unions in the same establishment. There may be some co-ordination in the bargaining strategies between the two groups with a settlement in the blue-collar area being conditional on the outcome of negotiations with the white-collar workers. It may also be true that employers find it difficult to resist white-collar demands that are similar to settlements just reached in the blue-collar field. If such added strength is derived from a direct association with a blue-collar union, it is important to note that virtually all the establishments in the survey reporting a collective agreement covering office employees also reported an agreement for non-office employees.

Even unorganized white-collar workers may be affected by unionism if employers who grant large increases to their blue-collar workers also feel compelled to grant similar increases to their white-collar workers. Employers may do this out of their sense of fairness or to prevent

discontent among white-collar workers caused by an erosion of traditional wage differentials. Instances of this type of behaviour abound in descriptions of employer wage-setting practices.[1] However, the extent of such union spillover effects is an open question. Although parallel increases for white- and blue-collar workers may occur occasionally, employers concerned with maximizing profits are not likely to follow this practice when the result would be wages for office employees that are far in excess of market rates. Besides, it is doubtful that manufacturing employers, in setting their white-collar wage policies, are yet to the point where threats of union organization must be taken into account.

In order to estimate the size of both the direct and spillover effects on white-collar wages, it is only necessary to add a dummy variable, CA_W, indicating the union status of the white-collar workers, to the previously used regression equations. With the dependent variable being a white-collar wage rate, the coefficient of CA_W will indicate the direct union effect while the coefficients of the (blue-collar) union impact variables will indicate the size of spillover effects. The set of control variables can be interpreted as before.

Actually, if the coefficients of the union impact variables are of roughly the same magnitude as reported in the previous chapters, two interpretations would be possible. First, it might signify that employers extend completely the wage increases granted to their unionized employees to the rest of their work force. Or, it might signify that both white- and blue-collar workers are subject to some common influences that are independent of unionism. Large union impact coefficients may mean that employers in concerntrated industries with relatively low wage costs may pay high wages to all their employees because they do not maximize profits and find it relatively easy to pursue a high wage policy. Our interpretation of these variables as indexing variations in union effects would be open to some doubt. The result would be uncertainty as to the extent to which industries pay high wages due to union effects or other causes.

On the other hand, if the union impact variables have little effect on white-collar wages, this would not only indicate that this type of spillover effect is small, but also would aid in the interpretation of the union impact variables. Presumably, if employers failed to maximize profits in their wage decisions, they would grant high wages to all their employees. Hence,

[1] See, for example, Rees and Schultz, Workers and Wages in an Urban Labor Market, p.45.

if white-collar wages are only weakly related to such variables as U, CR, and W/VA, this would raise our confidence in interpreting their effect on blue-collar wages as indicating variations in union impact.[2]

Variant I and II Model Results

In Table 20, the regression results for six white-collar occupations are presented. These occupations were selected for analysis so as to cover both male and female workers and to maximize the number of establishments

TABLE 20

VARIANT I AND II: REGRESSION RESULTS FOR WHITE-COLLAR OCCUPATIONS

OCCUPATIONS	PS	CA_W	CA	CR·CA(10^{-2})	W/VA·CA(10^{-2})	U·CA(10^{-2})	PS·CA	R^2	Number of Observations
Female:									
Bookkeeping Machine Operator	.0740 (5.4)	.1048 (1.7)		–	–	–	–	.188	724
	.0847 (5.7)	.1099 (1.8)	.0581 (1.8)					192	724
	0396 (1.6)	.0803 (1.3)	−.2217 (1.4)	.284 (1.2)	−.4415 (2.2)	.000 (0.0)	0702 (2.3)	.207	724
Junior Typist	.0547 (6.1)	.2433 (6.1)				–	–	.208	867
	.0663 (3.4)	.2517 (6.4)	−.0823 (3.4)	–	–		–	.218	867
	.0600 (3.5)	.2481 (6.2)	−.0592 (0.5)	.239 (1.4)	−.267 (1.7)	.098 (1.2)	−.0010 (0.0)	.227	867
Senior Secretary	.1487 (5.2)	.3173 (5.2)	–	–	–	–	–	.184	1272
	.1570 (10.2)	.3243 (5.3)	−.0054 (1.4)		–	–	–	.185	1272
	.1276 (5.1)	.3011 (4.8)	−.2211 (1.3)	.221 (0.8)	−.204 (0.8)	.094 (0.7)	.0382 (1.2)	.188	1272
Male:									
Senior Accounting Clerk	.0697 (2.9)	−.0163 (0.2)		–	–	–	–	.054	685
	0639 (2.5)	−.0224 (0.2)	.0379 (0.6)	–			–	054	685
	.0506 (1.1)	−.0332 (0.4)	.1544 (0.5)	−.0914 (0.2)	−.6065 (1.6)	−.0198 (0.1)	.0258 (0.5)	.058	685
Order Clerk	.0185 (0.8)	.0284 (0.3)				–	–	.054	774
	0437 (1.8)	.0384 (0.4)	−.1530 (2.8)			–		.063	774
	.0280 (0.7)	.0557 (0.6)	−.4964 (1.9)	4912 [illegible]	575 (1.4)	.072 (0.4)	0121 (0.2)	.069	774
Senior Clerk	.0466 (2.0)	−.1364 (1.6)						103	586
	0595 (2.4)	−.1226 (1.4)	− 1012 (1.5)					.107	586
	.0436 (1.1)	−.1432 (1.6)	−.2353 (0.8)	.0988 (0.2)	−.313 (0.8)	.394 (1.7)	−.0019 (0.0)	.116	586

[2] The interpretation of the union impact variables could also be clarified by examining their relationship with wage levels during a nonunion period. Unfortunately, data comparable to the kind used in this study are not available for such periods. Moreover, culling the available scraps of historical information to examine their consistency with our hypotheses is a major undertaking that we have viewed as beyond the scope of this study.

that could be included in the regressions. The survey actually reports weekly salaries for office workers and the regressions were run using the weekly rate as the dependent variable. To convert the results to an hourly basis, and thereby ease comparisons with production workers, the coefficients have been divided by 37.5, the predominant standard hours per week of office workers in manufacturing.[3]

The regressions indicate that the few female white-collar workers whose wages are determined by collective bargaining have a considerable wage advantage. But this is not the case for males. The union-nonunion wage differential for the three female occupations averages 10 per cent. In contrast, for two of the male occupations, the differential is not significantly different from zero and in the other case the differential is negative and just statistically significant.

There is no apparent reason for this difference in impact by sex. Perhaps it may arise because the male jobs are senior positions that do not necessarily come under the terms of the collective agreement. Alternatively, it may reflect the previously noted tendency for union-nonunion wage differentials to decline with skill level. All the male white-collar jobs have substantially higher earnings than the female jobs. However, this latter explanation is not totally convincing as within the group of three female occupations selected for analysis, the union-nonunion wage differential is highest for the highest paid occupation.

Contrary to expectations, having a union in the plant appears to depress the wages of the associated white-collar workers. In five out of the six occupations, the coefficient of the non-office collective agreement dummy variable, excluding all other union impact variables from the regression, is negative. Moreover, the negative coefficients, although generally small, are statistically significant in three cases. The meaning of this result is also not apparent. It clearly indicates that spillover effects as discussed above are not important on average. But it would seem implausible to suggest that a form of wage fund concept is operating with wage gains by unionized production workers being made, in part, at the expense of white-collar wages.

The negative values might stem from the omission of the union impact variables from the equations. In the truncated equations, some of the variation in wage rates that should be attributed to unionism may have been picked up by the control variables. Note, for example, how the coefficient of the plant size control variable falls with the addition of the

[3] Canada Department of Labour, Working Conditions in Canadian Industries, (Ottawa: Information Canada, 1970) p.137

union impact variables. However, when the overall blue-collar union effects were calculated using mean values for all the union impact variables, the reduction in the size of the negative effects proved to be small.[4] The reason for the small negative effects remains largely unexplained.

The most intriguing results derived from the regressions concern the apparent limited role that the (blue-collar) union impact variables play in explaining white-collar wages. Although the coefficients usually have the expected sign, they are typically quite small and statistically insignificant. The coefficient of CR·CA has the right sign in five out of six cases but is never statistically significant. The size of the coefficients are typically only one-half of those found in the comparable equation involving the basic male rate. The coefficient in the male order clerk equation is quite large but its standard error is also large. It should be observed as well that in this equation W/VA·CA has the wrong sign. In the other five equations the coefficient of W/VA·CA is negative as expected but only statistically significant in two cases. Again, the size of the coefficients are much smaller than in the production worker equations. Limited effects are particularly apparent in the case of the U·CA and PS·CA variables. For each, there is only one occupation where the coefficient is statistically significant with the right sign. In all other cases, the coefficients are small.

The unimpressive role of the union impact variables for white-collar workers is confirmed by the small increase in explained variance associated with the addition of these equations to the regression equations. The null hypothesis that all the coefficients of the union impact variables are equal to zero can only be rejected at the 95 per cent confidence level for two occupations – junior typist and bookkeeping machine operator. Even for these occupations, the values of the F-statistics are quite low.[5]

Summary

The findings reported in this Chapter differ from studies of the United States experience in that in Ontario the direct wage effects of unionism on female white-collar workers were found to be substantial. For female white-collar workers, the union-nonunion wage differential averaged ten per cent or about the same as for female basic labour, whereas for male

[4] The calculated overall effects, in the order of occupations appearing in Table 20 are –.02, –.07, –.02, +.05, –.12 and –.08.

[5] The junior typist and bookkeeping machine operator F-statistics are 4.2 and 3.3 respectively. This compares with a critical value of 3.0 (95 per cent confidence level).

white-collar workers, the differential was virtually zero. On the other hand, the results agree with the Raimon and Stoikov study[6] which also reported no spillover effects of blue-collar unionism on white-collar wages. The most intriguing result concerns the apparent minimal role that the union impact variables play in explaining white-collar wages. This aids in the interpretation of blue-collar regression results. Because no relations were found, it seems more likely that the union impact variables are actually capturing variations in union effectiveness rather than a more general cause, such as employer ability-to-pay, which would presumably affect the wages of white- and blue-collar workers alike.

[6]Raimon and Stoikov, "The Effects of Blue-Collar Unionism on White-Collar Earnings," p. 358-74.

CHAPTER VII

SUMMARY AND CONCLUSIONS

In this chapter, the results of this study are placed in perspective through an examination of the following topics. First, general limitations concerning econometric research and the particular limitations of this study are described. Second, the unique aspects of the model and data used are contrasted with the existing literature. Then, the main substantive findings are summarized and tied together. Where possible, these are related to the results of other studies. In a final section, some implications flowing from the analysis and directions for further research are discussed.

Limitations

Most attempts at measuring economic concepts through econometric techniques are fraught with difficulties and this study is certainly not an exception. Many choices have to be made on the variables to be included in the analysis and their exact specification. But perhaps the problems faced by this study in this regard are more critical than usual. The absence of a well-developed theory of union behaviour leads to uncertainty regarding whether all the important variables have been included. A similar specification problem arises due to a lack of concensus on the forces that would shape wage structures under nonunion conditions. Because of the focus on variations in union effects and spillovers, both rather novel topics, it has not been possible to build on the results of other studies of wage structures. Another problem is connected with our inability, due to a lack of data, to include in the wage determination equations all the variables that competitive labour market theory suggests are relevant. Our equations leave a substantial proportion of the variation in wage levels unexplained. Yet, one can not be certain if alternate specifications or the addition of more variables would materially affect the results.

This does not mean that attempts to measure the union impact on the wage structure are pointless. Despite their difficulties, such measurement attempts are bound to yield valuable insights and are vastly superior to the reliance on vague impressions. In general, econometric research must be viewed as a process where firm conclusions gradually emerge due to refinements made in a number of studies that explore alternate analytical approaches. As this study has been concerned with some previously

neglected topics, the findings must be put forth with appropriate caution.

Some limitations of the study also stem from the data base being restricted solely to manufacturing establishments in Ontario in 1969. For example, it is not clear if the union-nonunion wage differentials in Ontario are larger or smaller than similar differentials in other parts of the country where wages are low and unions not well-organized. On the one hand, it may be expected that unions in poorly unionized communities will be weak due, in part, to the ready availability of nonunion labour. But this might be outweighted by the inclination of unionized workers in poorly organized communities to draw comparisons with comparable unionized workers in higher wage areas of Canada. At the extreme, it is known that some unions attempt to reduce regional differentials or eliminate them entirely. Therefore, uncertainty exists whether union-nonunion wage differentials elsewhere in Canada are lower or higher than those observed in Ontario. However, given the concentration of Canada's manufacturing activity in Ontario, it would seem unlikely that a broader based study would yield substantially different results for this sector.

Restricting the analysis to the manufacturing sector may have biased the results downward. Ashenfelter's study,[1] the only one giving union-nonunion wage differentials by major industry division, shows larger differentials outside manufacturing, particularly in construction. One wonders, as well, if the results would be applicable to the rapidly expanding public and quasi-public sectors where the employer is not closely restrained by market conditions.

The suitability of the year 1969 for our cross-section analysis may be questioned. A number of writers have expressed the view that when actual rates of inflation exceed expected rates union-nonunion wage differentials will be reduced due to the fixed term of collective agreements. As there was a general acceleration in the rate of price inflation throughout the 1960's, reaching an annual rate of 4.6 per cent in 1969,[2] conceivably union-nonunion wage differentials may have been depressed during this year. The main empirical support for the counter-cyclical behaviour of these differentials rests on studies covering the depression and the immediate post-World War II period.[3] However, it is not obvious that the

[1] Ashenfelter, "Racial Discrimination and Trade Unionism," p. 449.

[2] Statistics Canada, Canadian Statistical Review, (Ottawa: Information Canada, various issues).

[3] Lewis, Unionism and Relative Wages, pp. 195-241 and Albert Rees, "Postwar Wage Determination in the Basic Steel Industry," American Economic Review, XLI, No. 3 (June 1951), pp. 380-404.

same behaviour would apply to the less extreme post-war business cycles. In fact, the only study examining the post-war experience was unable to discover any relationship between date and the size of union-nonunion wage differentials.[4] Given the slowly accelerating inflation of the 1960's, it seems unlikely that the differential would be greatly affected by the prevailing economic climate.

The Approach

It has been noted that prior studies using industry aggregated data led to estimates of the union-nonunion wage differential that are roughly double those reported by Lewis. But these estimates may be subject to an upward bias if it is true that wages in both union and nonunion establishments are positively correlated with the extent of industry unionism. As both theoretical considerations and some empirical evidence suggest that this will be the case, a disaggregated approach seems required.

The survey tape of the Canada Department of Labour, giving data at the establishment level, makes such an approach possible. As each establishment is allocated an S.I.C. code, industry as well as establishment variables can be included in the analysis. Moreover, the survey allows the use of occupational wage rates rather than average earnings so that problems connected with making adjustments for sex and skill levels do not arise. In short, the survey represents a valuable but hitherto unexploited data source that undoubtedly can be used to advantage in future studies of wage structures.

The model used in the analysis is distinguished by the fact that only two basic sets of factors are presumed to influence wages: (i) those consistent with competitive market influences and (ii) those related to unionism. The presumption is that in the absence of unions employers would maximize profits and labour markets would operate competitively. Unions are assumed to have a variable effect on wage rates determined by factors that, in the main, influence the elasticity of demand for union labour. For the first time, an extensive analysis is carried out using both the union status of the establishment and the extent of the industry unionism as explanatory variables. The model is also unique in that an attempt is made, using disaggregated data, to come to grips with the likely possibility that nonunion employers attempt to forestall unionization by raising their wage rates.

4 Reber, "The Effects of Unionism Upon Relative Wages," pp. 38-41.

The Substantive Results

First, consider what has been called the observed union-nonunion wage differential. After controlling for the effects of region, city location, working conditions (ratio of males and a durable-nondurable classification), employment growth, and plant size, it was found that union establishments do pay higher wages than others but that the differences are not as large as reported in some other studies. For male basic labour the wage differential is only nine per cent.

When union effects are allowed to vary, it is found that the extent of industry unionism and wages as a proportion of value-added are both important determinants of wage rate levels in union establishments. Although not as prominent, industry concentration and plant size also have positive effects on union wage rates. Profits appear to have a questionable role with the coefficients losing their statistical significance in the presence of other union impact variables. Of all the variables tried, only average establishment size in an industry, a crude index of barriers to entry, proved to play no role in explaining wage rate levels. With the exception of plant size, these effects were found to be quite consistent between all the occupations analyzed. Although both the male and female basic rates appeared related to the plant size union impact variables, this was not the case for the more narrowly defined occupations. For the skilled occupations, the industry concentration union impact variable played a weak role as well.

The strong positive association between the extent of industry unionism and union wage scales contradicts the conclusion of no relationship in the study by Bailey et al.[5] This is surprising as their study resembles our own as it also uses establishment data. Perhaps the difference arises because they use average hourly earnings as the wage rate variable and basically a two-digit rather than three-digit industry classification. In any event, our finding strongly suggests that estimates of union-nonunion wage differentials based on industry aggregated data are biased upwards. The study by Bailey et al also reported no relation between industry concentration and wage rates, whereas this study, along with others, shows such a relationship. Another implication flowing from the analysis is that the often observed relationship between wages and plant size may, in part, be the result of unionism, particularly for the unskilled.

[5] Bailey, King and Schwenk, "Wage Differentials," p. 397.

The extensive variation in union effects raises the possibility that the union-nonunion wage differentials in terms of establishments and employees might differ considerably. Using the regression results, the calculated differential in terms of employees for basic male labour was fourteen per cent.

Our interpretation of the union impact variables as indexing variations in union effects might be questioned. Others have included variables such as industry concentration and wages as a proportion of value-added in wage determination equations with the rationale that they reflect an employer's ability-to-pay.[6] The presumption is that departures from profit maximization would operate to raise the wages of union and nonunion workers alike. Conceivably, the union impact variables might be reflecting this type of influence rather than variations in union impact. However, it was found that for the predominately nonunion white-collar occupations analyzed, the union impact variables played virtually no role in explaining wage rate levels. Non-maximizing behaviour in the white-collar field appears limited and our interpretation of the union impact variables is supported. Unfortunately, it has not been possible to test further the interpretation through an examination of the relationship between production workers' wages and the union impact variables during a prior nonunion period.

Regardless of the interpretation, this study has demonstrated some rather strong empirical relationships. Industries with low labour costs, a concentrated product market and a high level of union organization do pay high wages within given blue-collar occupations. These relations, which are inconsistent with traditional justifications for wage differentials, must be kept in mind when the relative role of competitive labour market and institutional factors are being considered as determinants of the wage structure.

Among the control variables, it was found that the ratio of males has a surprisingly strong influence on wage rate levels. This raised the question whether a high proportion of females might more properly be viewed as a result, rather than a cause, of low wages. If this is the case, the inclusion of this variable in the wage determination equations would be wrong. Omitting the ratio of males from the male basic rate regressions has the effect of increasing the establishment union-nonunion wage differential only slightly to ten per cent. But the size of the union impact variables is

---

[6] Stuart O. Schweitzer, "Factors Determining the Inter-Industry Structure of Wages," Industrial and Labor Relations Review, Vol. 22 No. 2 (Jan. 1969), pp. 217-25.

increased substantially with the result that the differential in terms of employees rises to twenty-two per cent.

As union-nonunion wage differentials might vary among groups of production workers, they were calculated for a series of selected occupations. Consistent with a number of other studies, it was found that for basic labour the differential for females closely approximates that for males but that for the skilled occupations the differential, even in absolute terms, is considerably smaller than the differential for the unskilled and semi-skilled occupations. The skilled occupations probably represent less than one-quarter of all production workers eligible to join unions.[7] It would seem, then, that a global estimate for all production workers of the union-nonunion wage differential in terms of employees would only be slightly less than that previously indicated for the unskilled. Assuming that approximately one-quarter of all production workers are skilled and receive no wage advantage from unionism, while all the rest follow the pattern of the basic male rate, yields a global estimate of the differential of from ten to seventeen per cent depending upon whether a high ratio of females is viewed as a cause or effect of low wages.

Also, an attempt was made to assess the impact of unions on nonunion firms. Although an equation was specified where all the spillover effect parameters are identified, the estimates were implausible. The coefficients in virtually all the equations imply total spillover effects that are negative at mean values. A specification error appears to be present. However, nonunion wages do follow a pattern that suggests that employers do respond to the threat of organization by raising wage rates. Particularly for the unskilled it was found that:

(1) increases in the extent of industry unionism lead to increases in nonunion wage rates, except at low levels of industry unionism and other union impact variables,

(2) nonunion wages increase more rapidly with plant size at high levels of unionization than at low levels, and

(3) nonunion wages are high under circumstances that favour a high union impact.

---

[7]No statistics are available in Canada listing production worker employment by broad skill levels. However, employment statistics for the United States use the following categories for blue-collar workers: (i) craftsmen, foremen and kindred workers, (ii) operatives and kindred workers, and (iii) labourers. In manufacturing, the skilled group, including foremen, represents less than thirty per cent of blue-collar employment. See the U.S. Department of Labor, Handbook of Labor Statistics, 1971 (Washington D.C.: Government Printing Office, 1972), p. 58.

From our regressions it is difficult to say precisely how large the spillover effects are. Nonetheless, some indication can be obtained by comparing the expected nonunion wage rate under conditions that favour a low union impact and spillover effects with the expected wage for a comparable union establishment assuming mean values for the union impact variables. With the ratio of males included as a control variable, such a procedure suggests a union-nonunion wage differential in terms of employees of twenty-three per cent for basic male labour.

Concluding Remarks

The analysis strongly suggests that the overall observed union-nonunion wage differential for production workers in manufacturing lies in the range of 10 to 17 per cent. This is quite close to the range suggested by Lewis and a number of the recent studies using disaggregated data for the United States.[8] Two conclusions emerge:

(1) Studies using industry aggregated data are likely to yield estimates of the union-nonunion wage differential that are biased upwards due to a strong positive correlation between both union and nonunion wage levels, and industry unionism.

(2) The union impact in Canada does not appear to differ markedly from that of the United States perhaps reflecting the basic similarity between the industrial relations institutions of the two countries.

Given the attention paid to unions as wage-setting institutions, the apparent small size of the estimated differentials may appear to many observers as surprising. If these estimates are correct, two explanations are possible. First, the analysis indicates that there are wide variations in union impact on the wage structure so that misconceptions may arise if publicity is focused on those unions in a favourable economic position. Second, the estimates do not take into account the possibility that nonunion wages may be affected substantially by unionism. The findings do suggest that spillover effects may be important but it is difficult to indicate their overall size with any accuracy.

If it can be presumed that labour markets would approximate competitive conditions in the absence of unions, then an estimate can be made of the size of the misallocation of resources associated with the

[8] The two studies based on the Survey of Consumer Finances appear as exceptions. See supra, p. 29.

establishment of a positive union-nonunion wage differential. Based upon the assumption that the sole effect of unions is a relative wage advantage for their members of 15 per cent, calculations for the United States indicate a reduction in national output of 0.15 per cent.[9] This efficiency loss may not appear substantial. But if spillover effects are important, both the actual union-nonunion wage differential and workers affected could be considerably larger. As the efficiency losses increase quadratically with union impact, these losses could be considerably greater than previously reported.

This suggests that future research should focus on further attempts at measuring spillover effects. Using the same data base, it may be possible to get a better estimate of the size of spillovers by using a categorical approach to the specification of both the union impact and spillover variables. Based on their size and industry characteristics, establishments could be classified as subject to various levels of union impact and spillover effects. The category where spillover effects are presumed to be nil could then be used as the basis for calculating union-nonunion wage differentials. Such an approach would be less likely to yield implausible parameter estimates. Extending the analysis to include the largely nonunion service and trade sectors, and the addition of further control variables to capture compensating differentials might aid the analysis as well.

As a final word, it should be pointed out that if unions are viewed as an intervening force in the functioning of otherwise perfectly competitive labour markets, then their wage policies are bound to be associated with a misallocation of resources. However, it must be recognized that wage-setting is only one dimension of union activity. The emphasis given to wage determination in this study does not imply that other aspects of union activity are unimportant. In any comprehensive evaluation of industrial relations policies, the allocative efficiency loss associated with a positive union-nonunion wage differential, along with other costs in the form of work stoppages and productivity restrictions, must be weighed against the role collective bargaining plays in democratizing work relationships and providing protection against the vagaries of the market.

---

[9] Albert Rees, "The Effects of Unions on Resource Allocation," The Journal of Law and Economics, Vol. 6, No. 2 (Oct. 1963), pp. 71-72.

APPENDIX A

POTENTIAL SOURCES OF BIAS IN ESTIMATES OF THE UNION-NONUNION WAGE DIFFERENTIAL BASED ON INDUSTRY AGGREGATED DATA

Let W_i be the average wage in a given occupation for the ith industry and let W_i^u and W_i^n be the average wages for that occupation in the union and nonunion sectors of the industry. Also, let W_i^c be the average wage rate that would prevail if there were no unionism at all in the industry, but with unionism throughout the rest of the economy being held constant. In this case the actual union-nonunion differential for a given industry has been defined in relative terms to be

$$R_i^u = \frac{W_i^u - W_i^c}{W_i^c}$$

Similarly, the spillover effect can be defined as $R_i^n = \dfrac{W_i^n - W_i^c}{W_i^c}$.

Finally, define the relative difference between the actual wages of all unionized workers and the actual wages of all nonunion workers to be $M_i = \dfrac{W_i^u - W_i^n}{W_i^n}$. This has been called the observed union-nonunion wage differential. Corresponding to the above, $\overline{R}^u$, $\overline{R}^n$, $\overline{M}$ can be defined as economy wide averages. For example, $\overline{M} = \dfrac{\overline{W}^u - \overline{W}^n}{\overline{W}^n}$ where W^u is the average wage for all unionized workers in the economy and $\overline{W}^n$ is the average wage for all nonunionized workers in the economy.

When only industry aggregated data are available, one method that has been used in estimating the union-nonunion wage differential has been to regress an industry average wage rate against the extent of trade union organization in the industry as one of several independent variables. The following regression equation is illustrative of what is frequently done:

$$\log W_i = a + b\, U_i + Z_i \qquad (1)$$

where – $\log W_i$ is the logarithm of the average wage of all workers in the industry,

---

[1] The use of the geometric mean follows the practice established by Lewis. Although more complicated, the same results can be obtained using the arithmetic mean.

– U_i is the extent of trade union organization in the industry,

– Z_i is a residual error term.

To simplify the discussion, control variables are omitted. The problem is how to interpret the regression coefficient b. In particular, it is necessary to determine under what circumstances it would be legitimate to interpret b as an estimate of the union-nonunion wage differential.

The observed average wage rate in an industry may be taken to be the weighted geometric mean of the union and nonunion wages in the industry, i.e.,

$$W_i = W_i^{u\,U_i} \cdot W_i^{n\,(1-U_i)}$$

$$\log W_i = \log W_i^n + U_i \log W_i^u - U_i \log W_i^n$$

$$= \log W_i^n + U_i \log (1+M_i).$$

But $\log W_i^n = \log W_i^c + \log (1+R_i^n) = \log W_i^c + \log (1+\bar{R}^n) + \log \dfrac{(1+R_i^n)}{(1+\bar{R}^n)}$

and, $\log (1+M_i) = \log (1+\bar{M}) + \log \dfrac{(1+M_i)}{(1+\bar{M})}.$

By substitution,

$$\log W_i = \log W_i^c + \log (1+\bar{R}) + U_i \log (1+\bar{M}) + U_i \log \frac{(1+M_i)}{(1+\bar{M})} + \log \frac{(HR_i^n)}{(1+\bar{R}^n)}. \qquad (2)$$

Assume that $\log W_i^c = \log \bar{W}_c$, a constant, then (2) is in the same form as (1), with

$$a = \log \bar{W}_c + \log (1+\bar{R}^n)$$

$$b = \log (1+\bar{M}), \text{ and}$$

$$Z_i = U_i \log \frac{(1+M_i)}{(1+\bar{M})} + \log \frac{(1+R_i^n)}{(1+\bar{R}^n)}.$$

The expected value of Z_i is zero but it will be uncorrelated with U_i only if $(1+M_i)$ and $(1+R_i^n)$ are both uncorrelated with U_i. Accordingly, the least squares estimate of b will only be an unbiased estimate of $(1+\bar{M})$ if $(1+M_i)$ and $(1+R_i^n)$ are uncorrelated with U_i. However, there is some theoretical reason to believe that in fact a positive correlation will exist between these variables. Hence, b will generally be biased upwards as an estimate of log $(1+\bar{M})$.

On the other hand, it must be recalled that our prime interest lies not in $\overline{M}$ but in $\overline{R}^u$. If $\overline{R}^n > 0$, as our theory suggests, $\overline{R}^u > \overline{M}$ and b will be biased downwards as an estimate of $\log(1+\overline{R}^u)$ for this reason. Consequently, the overall direction of the bias in the estimate of $\overline{R}^u$ derived from the coefficient b in (1) is indeterminate.

APPENDIX B

THE RELATIONSHIP BETWEEN MONOPOLY PROFITS PER MAN-HOUR AND REPORTED PROFITS

The purpose of this appendix is to show the types of assumptions that have to be made in order to get a relationship between monopoly rents per production employee man-hour absorbed by labour and observable measures of profits. Following Rapping[1] let P be total monopoly profits for the establishment, and λ be the fraction of monopoly profits absorbed by the various factors of production when profits are not maximized. As it is plausible to assume that at least part of the monopoly profits go to stockholders, λ is likely to be less than one, that is $0 < \lambda < 1$. Let θ be the proportion of absorbed monopoly profits (λP) available to production labour. This proportion will also have a positive value less than one, that is, $0 < \theta < 1$. Labour's share of monopoly profits is defined as $P_L = \theta\lambda P$. The share of monopoly profits going to stockholders is, by assumption, $P_S = (1-\lambda)P$. Labour's share can also be expressed as $P_L = \theta \pi P_S$ where $\pi = \dfrac{\lambda}{1-\lambda}$.

If it is assumed that the opportunity cost of capital (R^*) may be expressed as a linear function of the measured rate of return, $R^* = \psi_0 + \psi_1 R$, the excess rate of return to stockholders in a given period is given by the expression: $R_S = R - R^* = (1 - \psi_1) R - \psi_0$. (1)
It seems reasonable to assume that $\psi_1 < 1$. Otherwise, high levels of observed rates of return would be associated with low excess rates of return. Multiplying (1) by equity assets per man-hour gives:

$$P_S/MH = (1 - \psi_1) R : A/MH - \psi_0 \; A/MH \cdot \tag{2}$$

This in turn can be used to derive an expression for labour's share of monopoly profits per man-hour.

$$P_L/MH = \theta\pi \; (1 - \psi_1) R \cdot A/MH - \theta\pi\psi_0 \; A/MH \tag{3}$$

In this expression $\theta\pi \; (1 - \psi_1)$ is positive but $\theta\pi\psi_0$ has an indeterminate sign.

The proportion of absorbed monopoly profits may not be a constant but rather it may vary with union power. If the extent of industry union-

[1] Rapping, "Monopoly Rents, Wage Rates and Union Wage Effectiveness," pp. 37-38.

ism indexes union power it can be assumed that $\theta = \delta_1 + \delta_2 U$ where $\delta_2 > 0$. Using this expression, (3) can be rewritten as:

$$P_L/MH = \delta_1 \pi(1-\psi_1)\, R \cdot A/MH - \delta_1 \pi\psi_0\, A/MH + \delta_2 \pi(1-\psi_1)\, U \cdot R \cdot A/MH - \delta_2 \pi\psi_0\, U \cdot A/MH. \tag{4}$$

The only unambiguous prediction concerning signs in (4) is that $\delta_2 (1-\psi_1)$ must be positive. However, some subsidiary relationships must hold. If $\delta_2\pi\psi_0$ is negative, this must be because ψ_0 is negative. If this is true, then $\delta_1\pi(1-\psi_1)$ and $\delta_1\pi\psi_0$ must have opposite signs, the particular signs depending on δ_1. On the other hand, if ψ_0 is positive, $\delta_1\pi(1-\psi_1)$ and $\delta_1\pi\psi_0$ must have the same sign.

APPENDIX C

NOTES ON DATA SOURCES

The variables used in the analysis relate to either establishment or industry characteristics. The establishment variables of region, city location, plant size, and existence of a collective agreement are derived directly from the Canada Department of Labour wage survey tape.

The establishment's SIC code provides the link for the industry variables. The 3,422 establishments in the sample are, in effect, assigned to 123 three digit industries. There were only minor departures from this scheme when a lack of an appropriate data breakdown in a secondary data source forced the use of an aggregate statistic involving two or three three-digit industries. Only sixteen small industries in manufacturing were deleted due to data not being available.

Statistics Canada reports, based on the Annual Census of Manufactures, were used as the source of information on industry employment growth, ratio of males to total employment, average establishment size, and the ratio of production worker wages and salaries as a proportion of value added. In some industries, the Ontario figures are not published due to confidentiality restrictions. In these cases, estimates were made from Canada-wide totals.

As mentioned above, the Herfindahl index based on the value of factory shipments for Canadian manufacturing industries in 1965 is the concentration measure used in this study. This index, recently published by the Canada Department of Consumer and Corporate Affairs,[1] is the sum of the squares of each firm size expressed as a percentage of total industry size. In compiling the statistics both the establishment and enterprise concepts were used. The latter was defined as all establishments in single manufacturing industry which are under common control. As the enterprise more closely approximates the relevant economic decision-making unit, the concentration index based on this concept was selected for use.

In measuring profits, the basic concept used is the rate of return after taxes on equity for corporations. The rates of return were calculated from data appearing in the Statistics Canada publication, Corporation Financial Statistics.[2] One of the specifications of the profit variable also involves

[1] Canada, Department of Consumer and Corporate Affairs, Concentration in the Manufacturing Industries of Canada (Ottawa: Information Canada, 1971).

[2] Catalogue No. 61-207.

equity assets per production worker man-hour (A/MH). This posed a problem as no single source presents the data required for calculating such a ratio and separate estimates of equity assets and production worker man-hours could not be used as the publications involved vary greatly in their coverage. A/MH had to be estimated indirectly as the product of two ratios: (1) equity assets divided by total wages and salaries and (2) total wages and salaries divided by production worker man-hours. The former ratio was obtained from Corporation Financial Statistics and the latter from the Census of Manufactures.

The most formidable data problem encountered was the absence in Canada of published estimates on the extent of unionism for three-digit industries. The Canada Department of Labour Occupational Wage Rate Survey could have been used to make such estimates. However, as the survey is limited to establishments with twenty or more employees and has variations in response between industries, an alternate approach appeared advisable.

The estimates used were based on unpublished data from the Ontario Department of Labour. The Research Branch maintains a virtually complete file of all collective agreements in the province. Each agreement is classified according to the type of bargaining unit (office, non-office, technical, etc.), industry, and the approximate number of employees covered. From this record, the total number of production workers covered by collective agreements by industry was tabulated as of January 1, 1969. These data were then related to the total number of production workers by industry as reported by the Census of Manufactures to produce the required estimates of the extent of unionism.

To guard against errors, the estimates derived in this manner were checked against the collective agreement coverage reported both in the Canada Department of Labour Occupational Wage Rate Survey tape and also the publication "Working Conditions in Canadian Industry 1969."[3] Where discrepancies existed, explanations were sought and adjustments made. In some cases, this required reference to basic establishment lists. The estimates produced in this manner are presented in the following table.

[3] Canada Department of Labour, Economics and Research Branch, Working Conditions in Canadian Industry 1969, Report No. 13 (Ottawa: Information Canada, 1970).

TABLE 21

ESTIMATES OF THE EXTENT OF COLLECTIVE AGREEMENT COVERAGE OF PRODUCTION WORKERS (U) IN PER CENT IN ONTARIO MANUFACTURING INDUSTRIES JANUARY 1, 1969.

SIC*	U	SIC	U	SIC	U	SIC	U
101	74	201	71	274	46	334	89
103	45	211	63	286	42	335	74
105	55	213	57	287	52	336	63
107	70	214	42	289	69	337	75
111	62	215	24	291	77	338	87
112	62	216	83	292	86	339	35
123	23	218	34	294	70	341	100
124	55	219	66	297	65	347	56
125	100	221	20	298	70	348	73
128	90	223	0	301	76	351	68
129	58	229	56	302	81	353	23
131	55	231	34	303	48	355	100
133	100	239	33	304	63	356	85
135	34	243	72	305	73	357	85
139	50	244	38	306	49	365	60
141	64	245	40	307	64	369	0
143	95	246	92	308	13	371	100
145	100	247	92	311	79	372	53
147	53	248	10	315	65	373	57
151	55	251	31	316	76	374	19
153	97	252	49	318	36	375	57
161	40	254	41	321	91	376	60
163	100	256	57	323	100	377	5
169	80	258	21	324	74	378	82
172	80	261	35	325	86	381	50
174	53	264	80	326	100	383	30
175	32	266	40	327	100	385	33
179	42	268	46	328	24	393	27
183	100	271	100	329	100	397	30
193	72	272	60	331	62	398	39
197	87	273	72	332	92		

* 1960 Standard Industrial Classification.

BIBLIOGRAPHY

Ashenfelter, Orley. "Racial Discrimination and Trade Unionism." The Journal of Political Economy, Vol. 80, No. 3, Part 1 (May/June, 1972).

____________, and Johnson, George E. "Unionism, Relative Wages and Labor Quality in U.S. Manufacturing Industries." Industrial Relations Section, Work Paper No. 9, Princeton University (May, 1969).

Bailey, William R., King, Harlan W., and Schwenk, Albert E. "Wage Differentials and Establishment Size, Union Status and Concentration." American Statistical Association, 1970 Proceedings of the Business and Economic Statistics Section. Washington, D.C., 1971.

Becker, Gary S. The Economics of Discrimination. 2nd Edition, Chicago: The University of Chicago Press, 1971.

Behman, S. "The Wage Determination Process in U.S. Manufacturing." Quarterly Journal of Economics, Vol. LXXXII, No. 1 (Feb., 1968).

Brown, David G. "Expected Ability to Pay and The Interindustry Wage Structure in Manufacturing." Industrial and Labour Relations Review, Vol. 16, No. 1 (Oct., 1962).

Canada Department of Consumer and Corporate Affairs. Concentration in the Manufacturing Industries of Canada. Ottawa: Queen's Printer, 1971.

Canada Department of Labour, Economics and Research Branch. Working Conditions in Canadian Industry, 1969. Report No. 13, Ottawa: Queen's Printer, 1970.

Canadian Labour Congress. "Submission by the Canadian Labour Congress to the Royal Commission on the Status of Women." Ottawa, 1968 (Mimeographed).

Doeringer, Peter B., and Piore, Michael J. Internal Labor Markets & Manpower Analysis. Lexington, Mass.: D.C. Heath and Co., 1971.

Dominion Bureau of Statistics. Manufacturing Industries of Canada Summary for Canada, 1968. Catalogue No. 31-203, Ottawa: Queen's Printer, 1970.

____________. Manufacturing Industries of Canada, Type of Organization and Size of Establishments, 1967. Catalogue No. 31-210, Ottawa: Information Canada, 1972.

———————. Annual Report of the Minister of Industry, Trade and Commerce under the Corporations and Labour Unions Returns Act. Part II, Ottawa: Queen's Printer, 1969.

———————. Canadian Statistical Review. Ottawa: Queen's Printer, various issues.

———————. Corporation Financial Statistics. Catalogue No. 61-207, Ottawa: Queen's Printer, 1965-68.

———————. 1960 Standard Industrial Classification. Catalogue No. 12-501. Ottawa: Queen's Printer, 1960.

Dunlop, John T. Wage Determination Under Trade Unions. New York: Augustus M. Kelley Publishers, 1950.

Friedman, Milton. Price Theory: A Provisional Text; Rev. ed., Chicago Aldine Publishing Co., 1962.

Fuchs, Victor R. The Service Economy. New York: The Columbia University Press, 1968.

Galbraith, John Kenneth. "The Class Struggle in the Industrial State." Paper presented at the National Tripartite Conference, Ottawa, Ont., Oct. 26-29, 1969.

Geoffroy, Renee, and Sainte-Marie, Paul. Attitude of Union Workers to Women in Industry. Studies of the Royal Commission on the Status of Women in Canada. Ottawa: Queen's Printer, 1971.

Hamermesh, Daniel S. "White-Collar Unions, Blue-Collar Unions and Wages in Manufacturing." Industrial and Labor Relations Review, Vol. 24, No. 2 (January, 1971).

Hicks, J.R. The Theory of Wages. 2nd ed. London: Macmillan, 1963

Hildebrand, C.H., and Delehanty, G.E. "Wage Levels and Differentials." Prosperity and Unemployment. Edited by R.A. Gordon and M.S. Gordon. New York: John Wiley & Sons, Inc., 1966.

Imel, B., and Helmberger, P. "Estimation of Structure Profit Relationships with Application to the Food Processing Sector." The American Economic Review, Vol. LXI, No. 4 (Sept., 1971).

Johnson, George E. "Wage Theory and Inter-Regional Variation." Industrial Relations, Vol. 6, No. 3 (May, 1967).

———————, and Youmans, Kenwood C. "Union Relative Wage Effects by Age and Education." Industrial and Labor Relations Review, Vol. 24, No. 2 (January, 1971).

Johnson, Harry G., and Mieszkowski, Peter. "The Effects of Unionization on the Distribution of Income: A General Equilibrium Approach." The Quarterly Journal of Economics, Vol. LXXXIV, No. 4 (Nov., 1970).

Kilpatrick, Robert W. "The Choice Among Alternate Measures of Industrial Concentration." The Review of Economics and Statistics, Vol. XLIX, No. 2 (May, 1967).

Kumar, Pradeep. "Differentials in Wage Rates of Unskilled Labour in Manufacturing Industries." Industrial and Labour Relations Review, Vol. 25, No. 4 (July, 1972).

Lester, Richard A. "A Range Theory of Wage Differentials." Industrial and Labour Relations Review, Vol. 5, No. 4 (July 1952).

Lewis, H.G. Unionism and Relative Wages in the United States. Chicago: University of Chicago Press, 1963.

Levinson, H.M. "Unionism, Concentration and Wage Changes: Toward a Unified Theory." Industrial and Labour Relations Review, Vol. 20, No. 2 (Jan., 1967).

______________. Determining Forces in Collective Bargaining. New York: John Wiley & Sons, Inc., 1966.

McConnell, Campbell R. "Institutional Economics and Trade Union Behaviour." Industrial and Labour Relations Review, Vol. 8, No. 3 (April, 1956).

MacDonald, Robert M. "An Evaluation of the Economic Analysis of Unionism." Industrial and Labour Relations Review, Vol. 19, No. 3 (April, 1966).

Mason, Sandra L. "Comparing Union and Non-Union Wages in Manufacturing." Monthly Labour Review, (May, 1971).

Masters, Stanley H. "An Interindustry Analysis of Wages and Plant Size." The Review of Economics and Statistics, Vol. LI, No. 3 (Aug., 1969).

Montague, J. Tait. Labour Markets in Canada: Processes and Institutions. Scarborough, Ontario: Prentice-Hall of Canada, 1970.

Oaxaca, Ronald L. "Male-Female Wage Differentials in Urban Labor Markets," Industrial Relations Section, Work Paper No. 23, Princeton University, n.d.

Ostry, Sylvia, and Zaidi, Mahamood A. Labour Economics in Canada. Vol. II of Labour Policy and Labour Economics in Canada. 2nd ed., Toronto: Macmillan of Canada, 1972.

Peitchinis, Stephen G. Canadian Labour Economics: An Introductory Analysis. Toronto: McGraw-Hill of Canada Ltd., 1970.

Pencavel, John H. "Wages, Specific Training and Labor Turnover in U.S. Manufacturing Industries." Institute for Mathematical Studies in the Social Sciences, Stanford University, June, 1971.

Raimon, Robert L. "The Indeterminateness of Wages of Semi-skilled Workers." Industrial and Labour Relations Review, Vol. 6, No. 2 (Jan., 1953).

______________, and Stoikov, Vladimir. "The Effect of Blue-Collar Unionism on White-Collar Earnings." Industrial and Labour Relations Review, Vol. XXII, No. 3 (April, 1969).

Rapping, Leonard A. "Monopoly Rents, Wage Rates and Union Wage Effectiveness." The Quarterly Review of Economics and Business, Vol. 7, No. 1 (Spring, 1967).

Reber, Stanley R. "The Effects of Unionism Upon Relative Wages." Unpublished Ph.D. dissertation, University of Chicago, 1970.

Rees, Albert. The Current State of Labour Economics. Paper presented at the conference of Canadian Labour Economists, Queen's University, Kingston, Ont., Feb. 26-27, 1971. Reprint No. 16. Kingston, Ont.: Industrial Relations Centre, Queen's University, 1971.

______________. The Economics of Trade Unions. Chicago: University of Chicago Press, 1962.

______________, and Shultz, George P. Workers and Wages in an Urban Labour Market. Chicago: University of Chicago Press, 1970.

______________. "Postwar Wage Determination in the Basic Steel Industry." American Economic Review, Vol. XLI, No. 3 (June, 1951).

______________. "The Effects of Unions on Resource Allocation." The Journal of Law and Economics, Vol. 6 (Oct., 1963).

Reynolds, Lloyd G. The Structure of Labour Markets. New York: Harper and Bros., 1951.

Rice, R.G. "Skill Earnings and the Growth of Wage Supplements." American Economic Review Proceedings, (May, 1966).

Roa, P., and Miller, R.L. Applied Econometrics. Belmont, Calif.: Wadsworth Publishing Inc., 1971.

Rosen, Sherwin. "On the Interindustry Wage and Salary Structure." Journal of Political Economy, Vol. 77, No. 2 (March/April, 1969).

______________. "Trade Union Power, Threat Effects, and the Extent of Organization." The Review of Economic Studies, Vol. XXXVI (2) No. 106 (April, 1969).

______________. "Unionism and the Occupational Wage Structure in the United States." International Economic Review, Vol. III, No. 2 (June, 1970).

Rosenbluth, G. "Wage Rates and the Allocation of Labour." The Canadian Journal of Economics, Vol. 1, No. 3 (Aug., 1968).

Ross, Arthur M. Trade Union Wage Policy. Berkeley: University of California Press, 1956.

Segal, Martin. "The Relation Between Union Wage Impact and Market Structure." Quarterly Journal of Economics, Vol. 78, (Feb., 1964).

Schweitzer, Stuart D. "Factors Determining the Interindustry Structure of Wages." Industrial and Labor Relations Review, Vol. 22, No. 2 (Jan., 1969).

Shister, Joseph. "The Logic of Union Growth." in Readings in Labour Economics and Industrial Relations. ed. by Joseph Shister. 2nd ed., New York: Lippincott, 1956.

Slichter, Sumner. "Notes on the Structure of Wages." Review of Economics and Statistics, Vol. 35 (Feb., 1950).

Stafford, F.P. "Concentration and Labor Earnings: Comment." American Economic Review, Vol. LVIII, No. 1 (March, 1968).

Stigler, George J. "Information in the Labour Market." Journal of Political Economy, Supplement (October, 1962).

Throop, A.W. "The Union-Nonunion Wage Differential and Cost-Push Inflation." The American Economic Review, Vol. LVIII, No. 1 (March, 1968).

U.S. Department of Labour. Handbook of Labour Statistics, 1971. Washington, D.C.: Government Printing Office, 1972.

Weiss, Leonard. "Concentration and Labor Earnings." American Economic Review, Vol. LVI, No. 1 (March, 1966).

______________. "Concentration and Labor Earnings: Reply." American Economics Review, Vol. LVIII, No. 1 (March, 1968).